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Susan Moloney consistently transformed our scarcely legible scrawl and multiple nested inserts and deletions into a typed version of what we really meant.

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1
THE NATURE OF THE STUDY

The passing of the Hospital Insurance and Diagnostic Services (HIDS) Act in July 1958 was a milestone in the evolution of health insurance in Canada. The legislation enabled the federal government to negotiate separate cost-sharing agreements with each province for provincially administered hospital insurance. Programs that met certain reasonably well specified objectives and criteria and that covered acute treatment and convalescent and chronic care in approved hospital facilities were eligible for the federal contribution. Constitutionally, health is clearly a matter of provincial jurisdiction, and the provinces were to take part in the federal scheme voluntarily; however the financial implications effectively deprived the provinces of any real opportunity to refuse to participate. Ontario joined the program on 1 January 1959.

The resulting federal-provincial hospital insurance agreements were the culmination of a long series of efforts to find solutions to the increasingly complex problems of financing hospital care. There had been a gradual replacement of direct patient payments by an assortment of voluntary, commercial, and government plans,[1] and the enactment of the HIDS Act sped up that trend.

The four main principles and conditions underlying the hospital insurance program are:

1 Comprehensiveness: all inpatient and outpatient services to which the

1 As far back as 1934, Newfoundland began a combined program of hospital construction and prepaid medical and hospital care in outlying areas, known as the Cottage Hospital Plan. Saskatchewan introduced the first province-wide hospital care prepayment plan with universal coverage in 1947. Two years later British Columbia introduced a universal coverage hospital insurance scheme. Alberta began grants-in-aid of municipal prepayment plans for hospital care in 1950. For a review of these developments see Royal Commission on Health Services (1964) and Taylor (1978).

residents of a province are entitled under provincial law are to be provided either without direct charge[2] or subject to authorized 'copayment' charges.

2 Universality: provinces must make insured services available to all residents of the province.

3 Accessibility: the provincial law must make 'provision for insured services in a manner that does not impede or preclude, either directly or indirectly, whether by charges or otherwise, reasonable access to insured services by persons entitled thereto and eligible therefor' (HIDS Act, Subsection 3(2) (b.1)).[3]

4 Portability: the provincial law must 'make provision for the payment of amounts to hospitals in respect of the cost of insured services, and the payment for insured services provided to residents of the province who are eligible therefor and entitled thereto by hospitals that are owned or operated by Canada or are situated outside the province' (HIDS Act, Subsection 3(2) (b)).

All the provincial hospital insurance plans are publicly administered, though this was not explicitly required under the HIDS Act.[4]

It is evident that the political philosophy underlying the emergence of this comprehensive, universal 'social insurance' scheme was that hospital

2 The provincial plan could have a premium system if the premiums were not related to specific services, i.e. did not constitute a charge payable for the use of hospital care by patients. The premiums were thus a form of tax. Three provinces have had authorized charges for hospital care from the very beginning of their provincial plans. Ontario had no authorized charges for hospital care until 1978. Each province also imposes rather modest per diem 'differential charges' for private and semi-private hospital rooms.

3 'Reasonable access to insured services' is not so much contingent on whether or not a provincial law imposes authorized charges as it is on the amount and manner in which they are applied to users of insured services. There was no indication in the legislation as to what might constitute a charge that impedes or precludes reasonable access to insured services. Access to insured services is also influenced by the physical supply of services and their proximity to insured persons. Obviously, availability might be viewed as the paramount factor in judging the question of reasonable access, but this condition in the HIDS Act was never intended to regulate a province's hospital bed supply nor to control the location of hospitals providing insured services. Local public and professional capital and operational planning are the principal influences in that regard.

4 Public administration was, on the other hand, explicitly required in the case of provincial medical insurance programs under the Medical Care Act of 1968.

services should be readily accessible to all on the basis of need rather than ability to pay. The most significant economic implication from an individual perspective is that it offers publicly funded prepayment of the costs of hospitalization and thus, with the exception of small and widely scattered direct charges to patients, the removal of private risk-bearing.[5] But has Ontario's scheme succeeded in providing equal access to hospital services?

The nature and scope of the study

The principal objective of this study is to answer the following questions empirically: in Ontario in 1974-5, who used and benefited from the province's hospital insurance and to what extent? In particular, how were use and benefits distributed across socioeconomic classes and were the characteristics of those classes determinants of that distribution?

The precise scope of the study is best elucidated in the context of broader evaluations of the objectives of public insurance. Barer, Evans, and Stoddart (1979) suggest two 'health insurance objectives' (p. 20) for public insurance, namely 'risk reduction' and 'wealth transfer.' We shall evaluate Ontario's hospital insurance program indirectly according to each of those objectives.

To measure the success of an insurance program in reducing risk, one needs information on the distribution of expected expenditures and private insurance coverage by socioeconomic class before the program was introduced. Our more restricted look at expenditures after universal insurance was introduced allows us to 'calculate the extent to which various groups [would] be worse off if the program were eliminated, ceteris paribus (Manga 1978, 13). It thus makes possible a snapshot of the risk-bearing role of the Ontario program.

At the broadest level, the wealth transfer objective could be evaluated through a differential or local fiscal incidence analysis,[6] which

5 It also increases the availability of care, since the insurance plans ensure a relatively stable source of income and financial support for hospitals, which, at the time of the HIDS Act, were experiencing considerable financial hardship (Taylor 1978).

6 For a good definition and description of what is entailed in these studies see Musgrave and Musgrave (1976). Such studies invariably confront numerous and complex statistical and conceptual problems, so

would measure the redistribution of wealth among various socioeconomic groups, when both the monetary value of the hospital services and the costs of financing the program are incorporated into wealth. Subsets of this type of analysis might look only at the change in tax incidence or only at the change in benefit incidence. The latter would entail an examination of changes in the distribution of hospital services (and costs) across socioeconomic groups resulting from the introduction of public insurance. This may be further disaggregated into separate examinations of the benefit incidence before and after the program began. Our study examines the utilization and cost incidence of the Ontario hospital insurance program for one year. This may be viewed as a benefit incidence study under the assumption that the benefits of the program to families are represented by the 'dollar value of [hospital] services paid by the program on behalf of the families' (Beck and Horne 1976, 76). The question of who pays for the program is not addressed, and therefore no effort is made to estimate 'net' winners and losers.

Thus, this study makes a cross-section (static) analysis of who used and benefited from the hospital insurance program in Ontario over a twelve-month period during 1974-5; to borrow the terminology of public finance once more, it is a benefit incidence analysis rather than an expenditure incidence analysis.[7] While a case may be made for including consumption externalities in the evaluation of benefits, we assume that the

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much so that some authorities are highly sceptical of their validity and merit (Bird and de Wulf 1973, Meerman 1978). Other equally renowned scholars believe them to be meaningful and relevant to government policy (Gillespie 1979). Stuart and Blair (1971) made a local fiscal incidence analysis of the Medicaid and Medicare programs in the United States. For other relevant literature pertinent to the United States see Keintz (1976); Feldstein, Friedman, and Luft (1972); and Wilensky and Holahan (1972). Studies in a Canadian context include Lindsay (1978), Morreale (1978), and Boulet and Henderson (1979).

7 For an excellent description of what is involved in the latter studies see McLure (1974). Expenditure incidence examines the effect of public expenditures on all private incomes. For example, we would have to look at how the implementation of hospital insurance plans affected nurses' incomes, doctors' incomes, the incomes of firms that supply the technology and materials to hospitals, and so on. Benefit incidence analysis, on the other hand, looks more at the narrower issue of who receives the services provided by public expenditures (i.e. the hospital insurance plan) and estimates the monetary value of the services provided to patients.

recipient of hospital care or the recipient's family is the sole beneficiary.

There are two serious difficulties that are likely to plague benefit incidence analyses of hospital insurance programs: (1) the absence of the requisite data base, particularly one that permits an allocation of hospital care to appropriately defined and described beneficiaries; and (2) conceptual and data problems relating to benefit measurement, specifically, information sufficient to place a monetary value on services provided to patients.

With respect to the first difficulty, we conducted a household interview survey to develop a micro-data base, including important socioeconomic, demographic, and income information on a sample of Ontario Health Insurance Plan families. The survey data were merged with hospital use records of the Ministry of Health to generate a data base sufficient to permit an incidence analysis. The establishment of the data base and related matters are discussed in Chapter 3.

When estimating the monetary benefits of hospital services from data on use, one is tempted to use readily available data on average cost per day or per case. The inadequacy of such measures is acknowledged even by their users. In order to convert our data on use into viable measures of benefit incidence, we use a number of cost weights, including diagnosis- and hospital-specific inpatient marginal costs of care, per case and per day, developed in Barer (1981b). Chapter 3 provides a brief description of the method.

Chapter 4 contains the heart of the benefit incidence results. Hospital use and benefits are reported by income class, age of head of household, and family composition, and tests are used to identify statistically significant differences in use or benefits across these groups.

Finally, recognizing that benefit incidence results imply nothing about causality, we devote a large part of this study to a causal analysis of hospital use. Before the introduction of comprehensive hospital insurance, many official and independent studies showed that hospital use was positively related to socioeconomic status and suggested that income was an important determinant of access to hospitals.[8] With the advent of public

8 By far the best source of Canadian evidence is the Canada Sickness Survey 1950-51, prepared jointly by the then Dominion Bureau of Statistics and the Department of National Health and Welfare. A cursory review of the literature on hospital use may be found in Chapter 5 of the present study.

hospital insurance, one may well ask whether or not the pro-rich use differential is eliminated. Do the poor indeed use more hospital services than the non-poor when other potential contributory factors are taken into consideration? Does income continue to be an important determinant of the use of hospital care?

Although we are unable to compare the relative importance of explanatory variables with and without insurance, the analysis in Chapter 5 does provide evidence on the role of these variables when there is universal public insurance and no direct charges (the standard economic price variables).[9] Various dependent variables (including the benefit measures noted above) are related to income and other socioeconomic and demographic variables. Here, too, serious difficulties had to be overcome. In this stage of the study, the difficulties were statistical and were related to the 'appropriate' estimation technique. The two-stage method adopted (Chapter 5) distinguishes between the probability of admission and the subsequent period of inpatient care. The appendix provides a detailed justification for this choice and an empirical comparison with other methods that have been used in past analyses or that were considered and rejected as unsuitable for the purposes of this analysis.

Finally, Chapter 6 summarizes the findings, discusses their implications for government policy, and makes suggestions for further studies and analyses of the consequences of public health insurance.

In short, this study addresses a set of relatively straightfoward questions: In Ontario during 1974-5, (1) what was the pattern of hospital use by socioeconomic class? (2) what was the pattern of benefits from the hospital insurance program by socioeconomic class? (3) in the absence of direct charges, what characteristics (if any) of different socioeconomic classes explain the variations (if any) across those classes in hospital use and benefits?

Unfortunately, only the questions are straightforward. We begin in the next chapter by reviewing the literature that has shed light in the past on questions (1) and (2).

9 The distinction between an incidence analysis and a determinant or use analysis is the one made in an earlier study that examined the same issues for Ontario's medical insurance plan (Manga 1978).

2
A REVIEW OF THE EMPIRICAL LITERATURE

Perhaps the greatest sources of scepticism about the attainment of equality in the use of health services as a consequence of free universal insurance are studies conducted mainly by medical sociologists. They point to a number of attitudinal and behavioural characteristics of the poor and to features of the delivery system that, many have argued, are more important in determining the use of care than financial variables. Some of the factors cited are lack of knowledge and education about the symptoms of illness, ignorance of the availability of health services, relationships between medical personnel and patients of low socioeconomic status, the complexity and impersonality of institutions, problems of communication, the middle-class bias of health professionals, and a lack of integration and co-ordination with other social services, as well as certain racial, ethnic, religious, and life-style characteristics. Reviews of this literature may be found in McKinlay (1972) and Anderson (1973b). There are also a number of extensive bibliographies on the correlates of use; see for example Aday and Eichhorn (1972), and Freeburg et al. (1979).

Many have argued that the mere removal of financial obstacles to health care will not greatly affect the traditionally lower use of care by the poor. Supplementary programs directed to the poor are necessary, they believe, including a variety of social services and educational programs. Furthermore, 'because of a drastic mismatch between medical organization and lower income life-styles, the extension of quality care to lower income groups requires a radical reorganization' of the health care delivery system (Strauss 1969, 146).

Clearly the reasons why people use hospital services are complex, many-sided, and not yet well understood. The vast literature on the correlates of use attests to the impossibility of addressing all aspects in one study, so our focus in this chapter is limited to an examination of

empirical studies in the United Kingdom, the United States, and Canada that have ventured to answer the same straightforward benefit incidence question that we have asked. That is, we limit our review of the literature to the studies that examine the pattern of hospital use or benefits by socioeconomic class under publicly funded hospital insurance. It should be noted that 'socioeconomic patterns of hospitalization' is a much broader issue, which includes patterns of use under all methods of financing, public and private. By focusing on public insurance settings, we have deliberately narrowed our literature review to the studies most directly relevant to the empirical setting for our own study.[1] We do, however, include for consideration a small number of empirical studies of publicly funded programs, which, while not examining explicitly benefit incidence under public hospital insurance, at least suggest or imply what the patterns are likely to be.

The hospital insurance component of the National Health Service (NHS) in the United Kingdom has very much the same principles as the hospital insurance plan in Ontario. The most important of these principles are unlimited access to all medically necessary hospital care free of direct charges and universal coverage of all residents on equal terms and conditions. This principle is not seriously vitiated by the existence of a relatively small number of private hospital beds that serve those who prefer and can pay for such care. Thus the benefit incidence consequences of that aspect of the NHS are particularly relevant to this study.

The U.S. programs are quite different in many important respects from the Canadian program: most notably they rely more on patient charges, and they are not universal. They are nevertheless of interest. Medicare and Medicaid, introduced in 1966, were in the forefront of the 'great society' programs designed to help the poor (see Davis and Schoen 1978), and they received the largest and most rapidly growing share of the public funds devoted to the social programs enacted during that period

1 The literature addressing the determinants of hospital use is reviewed briefly in Chapter 5 by way of providing guidance for our own analysis of use. Strictly speaking, any study of use that includes socioeconomic determinants is a possible source of information on the incidence of hospital use by socioeconomic class. That literature is not reviewed in this chapter because it derives predominantly from other than public insurance settings.

(Davis and Reynolds 1976, 391). Furthermore, the American studies are also of methodological interest to this study, because their data bases bear a significant resemblance to the information used in the present study.

A BRIEF OVERVIEW OF THE UK EXPERIENCE

There have been very few studies of the use of hospital services by various groups in the population of the United Kingdom. Furthermore, social class differences in the use of hospital care have received far less attention than social class differences in the use of medical services; this is true for Canada and the United States as well. In a comprehensive review of the literature on inequalities and the National Health Service, Townsend (1974, 1186) remarked:

> discussion of inequalities between regions and areas is too often sealed off from discussion of the underlying inequalities of class, income ... Neither social class nor income level features as a variable in analysis or even as a term so far as I can discern anywhere in the 182 pages of the Chief Medical Officer's report "On the State of the Public Health" for 1972, or for that matter in the reports of the previous 3 years.

Abel-Smith and Titmuss (1956) studied the distribution of hospital discharges in 1949 for men between the ages of 25 and 64 for different social classes in England and Wales. They concluded that the lower social classes were not well represented among teaching hospital discharges and only had average representation in regional general hospitals in relation to their mortality. Later, Titmuss (1968, 196) asserted:

> We have learnt from 15 years' experience of the Health Service that the higher income groups know how to make better use of the services; they tend to receive more specialist attention; occupy more of the beds in better equipped and staffed hospitals; receive more elective surgery; have better maternal care; and are more likely to get psychiatric help and psychotherapy than low-income groups - particularly the unskilled.

Similar views were expressed earlier by Butler and Bonham (1963) and by Abel-Smith and Townsend (1965). More recently, Townsend (1974) presented a considerable amount of evidence in the form of mortality and morbidity statistics in support of his thesis that health status differentials between the social classes in Britain are widening, not narrowing, under

the National Health Service.[2] In a national study of the elderly in institutions, he found that more persons from non-manual (i.e. socio-economic groups that include professionals, employers, managers, teachers, and so on) than from unskilled or partly skilled manual occupational backgrounds were in geriatric hospitals (as opposed to psychiatric hospitals), 'even when some attempt was made to standardize among patients by degree of incapacity, confusion, and lucidity, and were also in the better endowed hospitals within these sectors - defined by furnishings and shared spaces as well as staffing ratios'. He suggested that, at least to some extent, both clinical and administrative decisions respecting the allocation of resources appear to be influenced by the status of the institutions and the social class of their patients. He also found that the poorer working classes may stay longer in certain health institutions for reasons of a lack of housing in the community or personal lack of capital or income 'or because the institutions in which they live develop a functional need for their labour, their lack of demand upon a hard pressed medical and nursing staff or their value for teaching' (Townsend 1974, 1187).

Earlier, Hart (1971) argued that the availability of good medical and hospital care tends to vary inversely with the needs for care in the population served. The reason for this is that the best-trained and most able doctors tend to practise in upper- and middle-class areas, where mortality and morbidity are least prevalent. Hart suggested too that upper- and middle-class patients demand more health services than working-class patients and tend to be more critical of the care they receive, with the result that they are served better.

In a more recent study, Le Grand (1978) supports the conclusion that the professional and middle classes make better use of the Health Service than the poorer social groups. A notable feature of Le Grand's study is that he normalizes or adjusts the use of health care services by the various socioeconomic groups for the difference in health status among them.

2 The evidence presented by Townsend that health status differentials between the social classes in Britain are widening rather than narrowing is often interpreted to mean that access to medical and hospital care is correspondingly unequal. However, it must be noted that this inference is valid only if it is assumed that access to health care had a significant impact on health status.

Unfortunately, Le Grand's analysis and data encompass all health services and do not isolate hospital services. Nevertheless, his results throw doubt on the popular belief that the National Health Service benefits the poor proportionately more than the non-poor.

The overall conclusion of the authorities on the National Health Service, such as Abel-Smith, Titmuss, Hart, and Townsend, is, we believe, accurately reflected by Townsend's (1974, 1189) assertion that 'the right of the sick to free access to health care irrespective of class or income, remains to be firmly established.'

But there are contrary views and conclusions. For example, information on the use of hospitals by children under five years of age based on a follow-up study of a national random sample of all legitimate births in March 1946 suggested that poorer children, particularly those of unskilled manual workers, were both more likely to be admitted to a hospital and to have longer stays (Douglas and Blomfield 1958).

Carstairs and Patterson (1966) undertook a study of hospital admissions using Scottish hospital inpatient statistics for 1963 in order to see if there were any differences in use of hospital facilities by the five social classes.[3] Standardized discharge ratios were established that correct for variations in the age structure in each of the social classes. The principal findings (summarized in Table 1) are that for males and females in class 5 (unskilled occupations) the admission rates as measured by discharges were above the population average by 49 per cent and 56 per cent respectively. There is no clear linear relationship between the admission rate and social class for either males or females. For males, a positive class gradient is found between class 2 and class 5, with those in class 1 having a higher rate than those in classes 2 and 3. For females too, the admission rate for class 1 is higher than for classes 2 and 3 but lower than for classes 4 and 5. Carstairs and Patterson (1966, 2) explain the higher rate in class 1 in relation to classes 2 and 3 this way:

3 Most studies of social class in the United Kingdom are based on the General Register Office's Classification of Occupations, 1960. This scheme distinguishes five social classes: class 1, professional; class 2, intermediate; class 3, skilled occupations; class 4, partly skilled occupations; class 5, unskilled occupations. Class 3 occupations are subdivided into manual and non-manual. The five categories can be grouped into middle class (professional, intermediate, skilled non-manual) and working class (skilled manual, partly skilled, unskilled).

TABLE 1

Inpatient hospital use data for Scotland 1963 (all ages)

	Social class				
	1	2	3	4	5
Admission ratios[a]					
Males	92.5	84.1	91.7	105.4	148.8
Females	98.6	95.0	92.0	100.0	155.5
Mean stay in days - ratios[a]					
Males	77.5	88.2	98.7	103.0	112.1
Females	88.2	93.3	99.4	102.8	106.8
Bed-days per 1000 population					
Males	1250	1683	1528	2150	2604
Females	1042	1175	1101	1336	2153

a Figures are ratios of age-adjusted rates to average population rates
SOURCE: Derived from Carstairs and Patterson (1966, Table 2)

The higher rate for class 1 may be quite valid but it is suggested, as in the case of higher mortality figures for this group, that it may be due in some measure to upgrading of occupation, which is most likely to be present in data collection which is less rigidly controlled than the Census and which would result in an excess of persons appearing in class 1, in relation to the Census data.

The data in Table 1 also show that members of social classes 4 and 5 have longer stays and therefore more total days. Finally, in each social class males were, on average, greater users of inpatient care than females. The authors also examined admission rates, lengths of stay, and bed-days of hospitalization by social class and broad diagnostic group (such as malignant neoplasms, nervous system and sense organs, circulation, and so on). The results generally parallel those described above. The higher average length of stay for the poorer social classes (particularly social class 5) is only partly explained by differences in morbidity. Socio-economic factors such as poor housing, lack of income, and the unavailability and expense of home nursing were deemed to be important causes of

the relatively high average lengths of stay for patients in class 5 (Carstairs and Patterson 1966, 7). The data on use presented by Carstairs and Patterson suggest that partly skilled and unskilled persons are relatively high users of hospital services. However, there is no completely accurate measure that can be used to judge whether they are indeed receiving their 'fair share' in relation to their morbidity. Some proxy ratios of use to need have been postulated (Forster 1976, Dawkins 1976, Steward and Enterline 1961), but there is no empirical evidence using those measures.[4] This, unfortunately, is also true for Canada and the United States.

Rein (1969a) reviewed some data on the use of psychiatric hospitals in the United Kingdom. The interesting observation he made is that for almost all age groups there is clear evidence of higher rates of admission for lower social classes. But the relationship is rather unclear for the 65-and-over age group, a fact that prompted Rein (1969a,49) to comment:

> The data illustrate that demographic factors can affect the social class utilization in quite different ways ... it might appear that stress in the physical environments among younger age groups increases the need for hospital inpatient service for lower social classes [and that] increased psychological stress in the environment of the upper classes may over time increase their need for mental hospital services. But this interpretation is speculative. The data do show the relationship between class and use of in-hospital services for the mentally ill is neither simple nor self-evident.

In another review of the social class differences in the use of health services in the United Kingdom, Rein (1969b, 808) found that 'poorer children, particularly those of the unskilled manual workers, were both rather more likely to be admitted to a hospital and when admitted to stay

4 Ideally, measures of use should be related to measures of need. Empirical studies have rarely attempted to do so, mainly because of the lack of suitable data on health status or needs. Townsend (1974, 1186) reports that 'a Finnish study published in 1968 showed that the average number of consultations per 100 days with a physician was higher in the lowest than in the highest income group, but when consultations were standardized in respect of days of sickness, the trend was reversed. Moreover, the advantage of the relatively rich was shown for both the acute and chronic sick. The lower the income, the higher the morbidity and the lower the utilization of medical services in relation to morbidity.'

longer.'

While the evidence is mixed, it would appear that the lower social classes use hospitals more than any other group. This may indeed reflect nothing more than their higher rates of morbidity. The lower social classes are admitted more often and stay longer once admitted than those with higher status occupations in most cases, though this trend does not hold for all conditions or through all age and occupational groups in a consistently linear fashion. The lower social classes also seem to have higher rates of use of outpatient services as well as psychiatric hospital services, though in higher age groups the latter trend reverses itself with respect to social class.

THE U.S. EXPERIENCE: MEDICARE AND MEDICAID

What is generally known as Medicare in the United States stems from the social security amendments of 1965, which established two related health insurance programs for persons aged 65 and over and some under 65 who are disabled. Medicare does not cover care that is not 'reasonable and necessary' for the treatment of an illness or injury; nor does it cover custodial care.

Under Part A of Medicare, the basic compulsory hospital insurance plan (BHI), the beneficiary in 1980 was required to pay a $180 deductible ($40 in 1966) for the first sixty days of hospitalization and $45 ($10 in 1966) per diem for the sixty-first to the ninetieth day. After ninety days a patient can draw on a 'lifetime reserve' of sixty days, during which he or she must pay $90 for each reserve day used. There are also deductibles, co-insurance, and maximum days of coverage for post-hospital extended care or home health care, and deductibles and co-insurance are also charged for outpatient hospital diagnostic services.

Part B, the Voluntary Supplementary Insurance Plan (SMI), covers physicians' services and some related care. Subscribers pay a monthly premium of $8.70 (in 1980), which is estimated to pay for half the cost of the plan, as well as a deductible $60 (in 1980) and twenty per cent co-insurance for all allowable charges above the deductible. Allowable charges are those considered customary and reasonable for a given service in a particular area.

The financing of the BHI over and above the cost-sharing by the

beneficiaries is by taxes levied on employers, employees, and self-employed persons as well as by general funds from the Treasury. The latter pay for benefits to persons who are not beneficiaries of social security or of railroad retirement plans. The financing of SMI, over and above the cost-sharing and premiums paid by beneficiaries, is out of general revenues.

Medicaid (Medical Assistance Program) was enacted as an intergovernmental grant under Title XIX of the Social Security Act of 1965 in order to help the poor pay their medical costs. The reimbursement formula is based on state per capita incomes, permitting 50 per cent federal matching for high-income states and, in states with very low incomes, up to 83 per cent federal participation for a wide variety of covered expenses. The objective was clearly to create incentives for the poorer states to increase their expenditures on medical welfare, and thus Medicaid is perhaps the first time the federal government overtly attempted to reduce regional and inter-state differences in income and levels of health payments.

There is some evidence that Medicaid and Medicare have led to noteworthy increases in the use of medical and hospital services by the poor in relation to higher-income groups and that they have made hospitals and nursing homes much more accessible to the elderly of all income classes (Andersen et al. 1972, Pettingill 1972, Loewenstein 1971, Donabedian 1976a).

Loewenstein (1971) analysed data from household interview surveys of two national samples of persons 65 and over who were also recipients of social security benefits. Information was obtained about the use of medical and hospital services for the year before (1965) and the year after (1967) the introduction of Medicare. His analysis shows that there was an increase of slightly more than 10 per cent in admissions to short-term hospitals, and an increase of over 10 per cent in average length of stay during the survey period; these increases together resulted in an overall increase in patient-days of about 25 per cent. Also, the extent of change in hospital use varied by the demographic, social, and geographic characteristics of the respondents. The increase in hospital use was larger for the 75-and-over age group than for the 65-to-74 group, larger for blacks than whites, larger in the south than in other regions, and larger in non-metropolitan urban areas than elsewhere. While the difference between blacks and whites was reduced slightly, it was far from eliminated. In

addition to changes in the volume of hospital use, there was a disproportionate rise in admissions for surgery, which increased by over 30 per cent as compared to the 11 per cent increase for all admissions. As was true for all admissions, the increase for surgical admissions was greatest among those of advanced age.

In a more recent study, Wilson and White (1977) noted a marked increase in the number of low-income people receiving hospital care in 1973, in relation to 1964. While their analysis is based on a survey sample including persons not covered by Medicare and Medicaid, they do report sharp increases in hospital discharges for the poor (i.e. with family income of under $6000 in 1973), many of whom would have been covered by Medicaid or Medicare or both. During the same period there was practically no change in the hospitalization rates of non-poor people, the majority of whom were presumably not eligible for Medicaid or Medicare. Wilson and White (1977, 636) conclude, 'the data indicate that some of the gaps that existed in 1964 between the poor and the non-poor have been narrowed or eliminated, particularly in hospital and out-patient physician utilization.'

The higher use of hospital care by the poor in the Medicare/Medicaid period reported by Wilson and White is consistent with that reported earlier by Keintz (1976). Noting that 'the lowest family income group includes many persons, sixty-five years and over, whose hospital care expenses ... are covered by the Medicare program' (p. 67), she reports an admission rate for the poor (income of $0-1999) in 1970 over double that for the highest income class ($17,500 and over). In general, an inverse relationship between admissions and income is noted.

However, some shortcomings in the programs have been reported. In relation to higher-income groups, the poor are thought to be receiving lower-quality care (Davis 1976). The care they receive is discontinuous, episodic, fragmented, and impersonal. The poor are more inclined to obtain services from general practitioners than from specialists, in a hospital outpatient department rather than a physician's office, to wait substantially longer for services, to travel farther, and to wait longer to obtain appointments. Racial discrimination and variation in the availability of resources result in blacks receiving fewer services than whites (Davis and Reynolds 1976). Moreover, Medicaid does not provide health care services for all poor persons but only for those falling within certain welfare categories (e.g. blind, disabled, aged, and single-parent families).

For example, 'in 1974, an estimated 9 million persons with incomes below the poverty level, or about 35 percent of the poor, were ineligible for Medicaid' (Davis and Reynolds 1976, 394). Clearly, gains in use are not shared by all poor persons.

Finally, the patterns of hospital use by income class observed by Loewenstein, Wilson and White, and Keintz must be interpreted carefully. Their analyses do not consider adequately the differences in health needs among the different income groups. Comparisons between persons of equal health status may therefore be expected to reveal greater pro-rich differences than may be apparent from data not adjusted for health needs (Davis and Reynolds 1976, Donabedian 1976). Nevertheless, the evidence in these studies suggests that the disadvantages suffered by the blacks and the poor have been reduced by the introduction of Medicare and Medicaid.

The study by Davis and Reynolds (1976) is particularly noteworthy because it overcomes some of the shortcomings of the studies reviewed above. The authors used individual data from the 1969 Health Interview Survey to investigate the patterns of use of hospital and medical services among the elderly according to both health status and income. Health status is measured by restricted activity days and chronic conditions. As was mentioned before, it is difficult to standardize for health status in an examination of patterns of use among income classes. Data on morbidity are one source of the difficulty. There is also little agreement as to which measures are analytically most appropriate. The measures used by Davis and Reynolds were admittedly 'crude.'

Using a Tobit regression analysis, they found that after adjusting for other determinants of use, the number of hospital episodes (admissions) and hospital days increase uniformly with income irrespective of health status (as shown in Table 2). Significant racial differences in hospital care exist, particularly in the South. A comparison of elderly people with comparable health revealed that average length of stay increased directly with income and that the differences were not totally accounted for by a greater abundance of hospital beds in higher-income areas or by higher levels of education. On average, higher-income persons whose health status was judged to be 'average' for the elderly had 44 per cent more hospital days per capita than lower-income elderly persons.

It is important to note that Davis and Reynolds' study also suggests

TABLE 2

Average hospital use by the elderly by health status and family income, United States 1969

Family income	Health status Good	Average	Poor
	Hospital episodes		
Under $5,000	0.114	0.210	0.362
$ 5,000-9,999	0.140	0.250	0.427
10,000-14,999	0.159	0.285	0.472
15,000 and over	0.177	0.312	0.512
	Hospital days		
Under $5,000	2.31	4.21	7.21
$ 5,000-9,999	2.78	4.93	8.16
10,000-14,999	2.85	5.02	8.29
15,000 and over	3.52	6.06	9.77

SOURCE: Davis and Reynolds (1976, 414)

that a reduction in net direct prices has a positive impact on use of services. Persons receiving public assistance with coverage under both Medicaid and Medicare[5] received 30 to 40 per cent more services than other low-income persons not receiving public assistance when other determinants of use, such as health status, sex, age, race, and education, were held constant. Indeed, for the poor receiving public assistance, the use of services was commensurate with that of middle-income groups with similar health needs. For persons with exclusive coverage under Medicare, the imposition of uniform deductibles and co-insurance provisions resulted in significant disparities in use on the basis of income; i.e. it deterred the poor.

In summary, the American evidence (such as it is) suggests two things:

1 Medicare and Medicaid appear to have resulted in an increase in the use of hospital services by the beneficiaries of the program.

5 In such cases, the Medicaid program would pay the premium, deductible, and co-insurance required under Medicare.

2 Davis and Reynolds' (1976) evidence suggests that, despite this increase, there is still a pro-rich pattern of hospital utilization under these programs.

THE CANADIAN EXPERIENCE

Though the Hospital Insurance and Diagnostic Services Act was passed in 1957 and most of the provinces had implemented their provincial plans by 1958 or 1959, there are few empirical studies of the benefit incidence of these public programs.

In 1947 Saskatchewan introduced universal comprehensive publicly financed hospital insurance - the first province to do so. Two studies shed some light on the benefit incidence effects of this program. In a study focusing primarily on the universal medical insurance plan, Beck (1973) also reported hospital visits by various income groups. Data for the study were obtained from three sources. Registration files contained information on certain socio-demographic characteristics of each family, such as family size, marital status of the head of the family, age of family members, sex, and location of the family. The medical service records had a patient history file containing all physician and hospital services received by the patient under the insurance plan. The income data for the sampled families were drawn from the individual tax returns held by the Saskatchewan Treasury Branch. A family is defined to be either a self-supporting person of any age, together with his or her spouse or dependents under eighteen years of age, or a single person over this age (whether self-supporting or not); this is the definition of family used by the Saskatchewan Medical Care Commission. The period covered in the study consisted of the calendar years 1963-8. A random sample of 40,000 families was selected in each year. The representativeness of the family size and income distributions of the samples was tested and found to be satisfactory.

The definition and measurement of 'accessibility' used in the study is the proportion of families of a given economic class that have not had medical services in a given year. The main relationship considered was the one between accessibility and income class over the six-year period. Families were grouped into seven income classes: no income, $1-1499, $1500-2499, $2500-4999, $5000-9999, $10,000-14,999, and $15,000 and over.

The medical service variables used were the whole medical care system, general practitioners' services, specialists' services, complete examinations, regional examinations, laboratory testing, home and emergency visits, hospital visits, major surgery, and minor surgery.

The data show that 'with the exception of hospital visits, a clear pattern of differential levels of use by income class is evident for all service types. Accessibility to services of the medical care system appears to vary directly with income class - the higher the income level the greater is the contact with physicians' (Beck 1973, 349). For hospital visits, he found that 'only three income classes display a general increase in contact with hospitals over time - the zero, the $1-1499, and the $2500-4999 income classes.[6] The remaining classes show a slight tendency to have less contact with hospitals over time. However, the magnitude of change is too small to be conclusive' (Beck 1973, 353). Furthermore, Beck's unconventional measure of use does not permit one to convert hospital visits by income class into hospital volume measures, such as patient days or monetary measures.

A later study by Beck and Horne (1976) provides further indirect evidence relevant to the interests of the present study. Using essentially the same data sources as the earlier study, they measured the benefits of the medical insurance program to families in terms of the dollar value of physicians' services paid for by the program on behalf of the families. Their results are presented summarily in Table 3. As can be seen, mean benefits increase with family income for services provided by general practitioners and specialists. The same pattern prevails for distribution by type of services, such as complete examinations, regional examinations, home and emergency visits, and major surgery. The only exception is hospital visits, the <u>medical</u> benefits of which are inversely related to family income. Beck and Horne (1976, 76) conclude that 'evidently members of poor families spend more time in a hospital (though not for major surgery) than do members of middle- or upper-class families. While there are doubtless many complex factors at play here, suffice it to say it is one of the few results in this study that is fully consistent with conventionally

6 Our reading of Beck's Figure 10 (p. 352) suggests that it is the $1500-2499 rather than the $2500-4999 income class that increased their hospital contacts. In other words, the three lowest income groups increased their contact rate with hospital care.

TABLE 3
Mean dollar value of medical benefits per family by income class for various types of services, 1967 sample data

Type of service	\$1-2499	\$2500-4999	\$5000-9999	\$10,000-14,999	\$15,000 and over	All classes
General practitioner	38.70	45.53	57.51	58.98	57.73	49.46
Specialist	22.71	26.89	40.28	44.80	60.83	33.10
Complete examination	5.62	6.31	8.43	9.55	10.15	7.25
Regional examination	14.30	17.07	22.04	23.17	23.47	18.82
Home and emergency visits	4.25	5.06	6.65	7.31	7.50	5.70
Hospital visits	7.85	7.15	6.84	6.31	5.75	7.08
Major surgery	15.76	18.67	24.93	27.28	36.13	21.44
All services	60.77	71.70	96.44	103.59	117.99	81.67

SOURCE: Compiled from Beck and Horne (1976, Tables 2, 3, 4, pp. 79-81)

held notions concerning the distribution of medical need in the population.'[7]

A more recent comprehensive look at the Saskatchewan experience with co-payment for hospital services (Horne 1978) also provides some empirical evidence pertinent to this study. Horne undertakes a two-part analysis of the determinants of hospital use for nearly 40,000 Saskatchewan families over the period 1966-71. While a detailed review of his results must wait until Chapter 5 (where they are compared with those of our analysis), a few results are worth noting here:

- The probability of at least one hospital admission per family was found to be negatively related to income level (indicating a greater likelihood of hospital contact for low-income families) and positively related to age of head of household and family size.
- For the families with at least one admission, the number of days and admissions showed the same general pattern as probability of admission.

Much more direct evidence of interest to the present analysis is found in the Statistics Canada (1977) study on the distributional effects of health and education benefits in Canada. The report is based on the data collected from a supplementary questionnaire to the Survey of Consumer Finances of 1975, which included questions on the use of educational, medical, and hospital services. We limit our remarks here to the findings most relevant to hospital use.

The data on hospital use are derived from answers by each person (or proxy respondent) to the following questions about the use of hospital services: the number of admissions to hospitals, whether there was surgery (a yes or no answer), and the number of days in hospital. The data were to cover the whole of 1974; thus the respondent was expected to recall his or her or someone else's hospital use for the entire twelve months. The survey itself was conducted in April 1975.

Persons who were admitted to hospital during 1974 were divided into two groups: (1) those who underwent no surgery while in hospital, and

7 Complex factors aside, the inference drawn by the authors 'that members of poor families spend more time in a hospital' is legitimate only under the assumption that medical benefits per inpatient hospital day are invariant across income classes.

TABLE 4
Percentage distribution of family units by income class, Canada 1974

Family income group	Unattached individuals	Families	All families and unattached individuals
$ 0- 4,999	51.1	10.6	21.7
5,000- 9,999	30.5	20.7	23.4
10,000-14,999	13.3	27.0	23.3
15,000-19,999	-	21.0	16.2
20,000-24,999	5.2	10.8	8.0
25,000+	-	9.8	7.4

SOURCE: Abstracted from Statistics Canada (1977, 12, Table C)

(2) those who received some surgical treatment. This differentiation was used to impute different costs of hospitalization to the two groups. Individuals in each province were assigned an average hospital operating cost per inpatient day depending on the type of treatment received. For cases without surgery, a standard per diem (operating expenditure per patient day) was the basic measure of benefits. Surgical cases had a sum based on fee-for-service payments to the physician(s) involved added to the hospital per diem. The sum was established as a weighted average of the chargeable services from the appropriate fee schedules. The overall average cost of hospital visits was about $102 per day.

The total allocated cost of hospital use was $3254 million, which was 75.6 per cent of the actual total cost of $4305 million in 1974. The total number of days estimated by the study was 32 million, which was 75.4 per cent of the actual total of 42.5 million days. The shortfall can be attributed principally to under-reporting due to lapses of memory.

The Statistics Canada estimates for hospital costs ignore the fact that surgical and non-surgical patients are likely to make different demands on in-hospital resources. In addition, the allocation of physicians' charges only to surgical patients imputes an unknown, but definite bias to any results deriving from the analysis. While drawing attention to the surgical/non-surgical distinction is not without merit, it is not clear that the Statistics Canada benefit imputation method is appropriate.

It is apparent from the foregoing that the Statistics Canada study had

TABLE 5
Percentage distribution of recipient units by family characteristics within income group, Canada 1974

Family income group	Unattached individuals	Married couples only	Families, married couples with children	All other families
$ 0- 4,999	60.3	15.8	10.1	13.8
5,000- 9,999	30.5	25.3	33.1	11.1
10,000-14,999	11.9	21.3	60.4	6.4
15,000-19,999	4.9	20.9	69.9	4.3
20,000-24,999	1.8	21.1	73.8	3.3
25,000+	2.9	15.7	78.5	2.9
Total	23.1	20.5	48.3	8.1

SOURCE: Abstracted from Statistics Canada (1977, 24, Table G)

to ask very simple questions about hospital use - questions that produce crude or highly aggregative data. The report admits to having encountered lapses of memory among the respondents, as the period to be recalled was from four to sixteen months. This problem is compounded by proxy responses for persons who were absent at the time of the interview.

The basic family characteristics used by Statistics Canada to examine the data on use were the family's income, the composition of the family unit, and the age and sex of the persons hospitalized. The percentage distribution of income by composition is given in Table 4. The large proportion of unattached individuals in the under $5000 income group reflects the tendency for this group to be composed of either the fairly young who have just left the family home or the elderly widowed who are living on their pensions.

Table 5 gives a breakdown of family characteristics of beneficiaries within income groups. The distribution shows that single persons tend to dominate the lower-income groups while married couples with children dominate the higher-income groups. Table 6 shows the distribution of all hospitalized persons by age and family income class. The most noteworthy (although not surprising) feature of the data presented there is the much greater proportion of low-income ($0-4999) hospitalized persons in the 65-and-over age group in comparison to all other income classes. Taken

TABLE 6
Percentage distribution of persons hospitalized by age of person within family income group, Canada 1974

Family income group	Under 15	15-24	25-44	45-64	65 and over
$ 0- 4,999	14.8	12.5	14.3	24.0	34.4
5,000- 9,999	25.0	15.4	21.8	22.0	15.9
10,000-14,999	32.8	13.0	31.6	17.7	4.9
15,000-19,999	28.5	14.1	34.7	19.2	3.5
20,000-24,999	22.7	12.7	37.9	23.5	3.2
25,000+	20.8	17.9	30.1	24.0	7.2
Total	25.5	14.0	27.6	21.0	12.0

SOURCE: Abstracted from Statistics Canada (1977, 41, Table M)

together, these three tables illustrate some important correlates of the use of health care. In the $0-4999 income range, 60.3 per cent of the family units are classed as unattached individuals, most of whom are 65 years of age or over and who are likely to need more health services than the young. On the other hand, married couples with children dominate the four income groups that cover the middle-income ($10,000-$24,999) and upper-income ($25,000 +) classes. The two main groups that use hospitals in this income range are children and persons 25 to 44. One reason that the latter age group constitutes the largest user of hospital care in the $15,000-24,999 and upper-income ($25,000 +) groups is that it has the greatest proportion of married women of childbearing age.

Table 7 presents the average number of days' stay per person by age, sex, and family income for those hospitalized. It is evident that low-income persons incurred relatively more days of hospital care. For females there is a reverse J-shaped relationship between family income and days' stay with the greatest number of days' stay for those in the lowest income bracket and the shortest stays for those in the $15,000 to $19,999 income bracket. This is true for males as well, though the average number of days' stay for men in the $25,000 plus income group is a little lower than for those in the $20,000 to $24,999 income group. This reverse J picture holds roughly for most of the specific age/sex groups, although the bottom of the J does shift one class for some groups. For the families

TABLE 7

Average number of days hospitalized per person by sex, age, and family income class for Canada (for persons with one or more discharges in 1974)

Family income ($)	Age										
	0-14	15-24		25-44		45-64		65+		All ages	
	Male and female	Male	Female	Male	Female	Male	Female	Male	Female	Male	Female
$ 0- 4,999	10.5	19.1	8.3	25.7	10.6	20.4	24.8	20.8	17.8	21.2	16.8
5,000- 9,999	8.7	5.5	9.6	15.9	9.7	18.7	16.1	24.3	17.9	18.1	12.5
10,000-14,999	9.2	8.6	8.4	8.4	8.5	19.9	14.8	26.5	16.9	14.4	10.2
15,000-19,999	7.5	5.9	8.1	9.0	8.3	12.2	12.5	17.8	12.1	10.6	9.2
20,000-24,999	5.5	5.4	6.4	16.4	7.3	15.1	13.5	7.6	23.6	12.7	9.3
25,000+	5.9	8.8	6.2	8.6	7.8	16.4	11.8	13.7	28.5	11.7	10.7
Average	8.4	8.4	8.4	12.4	8.6	17.6	16.5	21.7	18.2	15.6	11.7

SOURCE: Derived from Statistics Canada, catalogue 13-561 occasional, pp. 122-4, Table 12

TABLE 8
Average health and hospital benefits of recipient units, Canada 1974

Family income group		Unattached individuals	Families	Families and unattached individuals
$ 0- 4,999	T	$469	$ 934	$ 653
	H	390	777	543
5,000- 9,999	T	236	799	628
	H	196	665	522
10,000-14,999	T	150	646	589
	H	125	537	490
15,000-19,999	T		525	503
	H		436	418
20,000-24,999	T	106	509	501
	H	88	423	417
25,000+	T		501	493
	H		417	410

NOTE: T = total medical and hospital benefits; H = hospital benefits only
SOURCE: Derived from Statistics Canada (1977, 15, Table B)

in the $0-4999 income class, there is a pronounced sex differential in the 15 to 24 and 25 to 44 age groups: men are hospitalized over twice as much as women, and the average hospitalization for men from all income classes in the 25-44 group is 45 per cent higher than that for women.[8] In the 65-and-over age group, women in the highest income class used hospitals more than their lower-income counterparts; in this they were

8 These differentials are at least partially unexpected. Hospitalization related to pregnancy should result in a higher rate for women than for men. Therefore the overall averages for the 25-44 and 15-24 groups are indeed puzzling. In fact, they are 'directly contradicted by higher quality population data contained in, for example, Table 18(a) of the Supplementary Statistical Compendium of the HIDS Program (Health and Welfare Canada, June 1977) showing that in 1973 the ratio of female to male utilization rates in each of these age groups was approximately 2:1.' (We are grateful to an anonymous referee of our study for pointing out this anomalous finding to us). The specific results for the $0-4999 and $5000-9999 income classes may be explained at least partially by Table 4. It seems clear that a large number of the women in those income classes are single, which negates much of the influence of pregnancy on the incidence of hospitalization.

TABLE 9
Consumption of health services covered by medical care and hospital insurance programs: average per individual by income class of family in relation to average for Canada, 1974

Income class	Benefits received under: Medical care program: Outpatient	Total	Hospital insurance program	Both programs
Less than $5,000	1.41	1.66	1.94	1.85
$ 5,000-10,999	1.03	1.12	1.23	1.20
11,000-14,999	0.94	0.92	0.90	0.90
15,000-19,999	0.79	0.74	0.68	0.70
20,000-24,999	0.89	0.77	0.64	0.68
25,000+	0.81	0.72	0.61	0.64
Average for Canada	1.00	1.00	1.00	1.00

SOURCE: Based on data from Statistics Canada (1975 Survey of Consumer Finances) and estimates by Boulet and Henderson (1979)

unlike every other age group in that income class. Hospital use by women increases with age in all income classes except the lowest. The generally positive relationship between age and hospital use is also observed for men with family incomes of $5,000-19,999. The patterns for males in the lowest and upper income classes are quite different. Overall, males were hospitalized for longer than females, though Statistics Canada paradoxically reports that 'females used more hospital services than males, irrespective of their family's economic status' (Statistics Canada 1977, 42).

The average health benefits received by unattached individuals and families are presented in Table 8. The top figure in each cell is the estimate of total health (i.e. medical and hospital) benefits, while the bottom figure is the estimate of hospital benefits. It is evident that total health benefits and hospital benefits are inversely related to income for unattached individuals, families, and the total sample.

The pro-poor hospital benefit pattern is emphasized in a subsequent study by Boulet and Henderson (1979). Since both their data and the definitions and benefit measures are identical to the Statistics Canada study just reviewed, their findings are hardly surprising. Their results are presented in Tables 9 and 10. The authors conclude that the data

TABLE 10
Distribution of health care benefits received under medical care and hospital insurance programs across individuals ordered by income of their family unit and divided into quintiles, Canada 1974

Benefits received under	Quintiles[a]					
	First	Second	Third	Fourth	Fifth	Total
Medical care program						
Outpatient	25.2	20.3	18.7	19.0	16.9	100.0
Total	28.8	22.0	18.0	16.4	14.8	100.0
Hospital insurance program	33.1	23.9	17.2	13.4	12.4	100.0
Both programs	31.7	23.3	17.4	14.4	13.2	100.0

a Data for individuals are divided into five quintiles, each representing 20 per cent of the total. The first quintile comprises members of the families with the lowest total income, and the fifth quintile the highest total income

SOURCE: Boulet and Henderson (1979, 21)

'clearly reveal that as an individual's standard of living increases, his consumption of health services decreases. This observation on all health expenditures applies without exception, regardless of the level of income analyzed. More significantly, we find that this relationship applies even more to hospital care than to medical care' (p. 19). It can be seen from Table 10 that this pattern still holds true when presented in terms of quintiles rather than income brackets.

CONCLUDING REMARKS

Our brief review of studies of hospital use in the United Kingdom, the United States, and Canada suggests that under public hospital insurance schemes the poor are greater users and beneficiaries than the non-poor. The evidence for the United Kingdom is admittedly mixed, and no clear consensus can be claimed; that for the United States is considerably clearer. Medicare and Medicaid have made hospital services more accessible to the poor, although the studies invariably point out certain shortcomings of the programs that need reform. The Canadian studies, while few in number suggest that the low-income groups are the greater beneficiaries of the public hospital insurance program. Indeed, unlike the American and British studies, no Canadian empirical study that has been

published demonstrates or even suggests an opposite conclusion. In a later chapter we will compare our findings for Ontario with some of the studies cited in this chapter.

3
MEASURING THE DISTRIBUTION OF HOSPITAL COSTS

In the previous chapter we reviewed the rather limited empirical literature that has examined the pattern of hospital use by socioeconomic class under public insurance programs in the United Kingdom, the United States, and Canada. Particularly noteworthy were the relatively crude measures that have been employed as proxies for hospital use or benefits. Indeed, not a single study used any monetary estimates of hospital benefits for the UK or the United States. The Statistics Canada (1977) and Boulet and Henderson (1979) studies adopted straight per diems adjusted for surgical cases, a procedure that all authors admitted was a crude approximation to the measurement of hospital benefits. This chapter describes the various data bases and methods that we used to estimate hospital benefits. In addition, we describe data sources for the variables employed in the utilization analysis in Chapter 5.

THE OHIP FAMILY - HOSPITAL USE

In attempting to discover who uses hospitals in Ontario and to what extent, it is essential to define the consuming unit that is to be the subject of the analysis and to be able to categorize each unit according to certain socio-economic and demographic characteristics. The unit of observation upon which our analysis is based is the OHIP family, although selected use and benefit incidence results will also be reported on a person-specific basis.[1]

In Ontario, all records of hospital use by individuals or families are maintained by the Ontario Health Insurance Plan (OHIP). Upon

1 See Manga (1978) for a more detailed discussion of the OHIP family and the reason for employing it in an analysis of this nature.

discharging a patient, each hospital is required to complete a separation record giving details on diagnosis, length of stay, age and sex of the patient, and, of course, the patient's family OHIP number. The 'OHIP family' is composed of husband, wife, and unmarried children under 21 years of age who are not employed fulltime and who are dependent for support upon the insured person. Also eligible as dependents are children aged 21 and over who are financially dependent upon the insured person because of physical or mental infirmity, provided they were financially dependent before reaching age 21. For each such family, one family member (usually the household head) is designated as the OHIP registrant and all additional family members fall under that same identification number.

As in all studies such as ours, the choice of the population unit of analysis is dictated by data, policy, and procedural considerations. Given the peculiar definition of the family for health insurance purposes in Ontario and the fact that the requisite data are organized around this definition, our selection of the OHIP family as the most suitable unit of analysis was dictated in part by data and procedural matters. A study using an economic or census definition of the family would pose many complicated problems of data linkage.

From a standpoint of government policy it can be argued that the family is a more suitable unit of analysis than the individual for the present study. While it is literally true that the individual is the beneficiary of hospital services received, the externalities to the rest of the family are obvious. From a financial perspective it is the family that benefits from not having to pay hospital bills for its members. Moreover, as pointed out by Litman (1974, 495), there is much evidence that the 'family does indeed constitute a basic unit in health and medical care, exhibiting characteristic patterns of morbidity, response to symptoms and utilization of medical services and facilities'.

It is not an uncommon practice, as illustrated in Chapter 2, to determine the hospital use by families or individuals by means of a household survey. However, surveys are inferior to OHIP data for a number of reasons:

- because memories fade with time, the period for which respondents can be expected to recollect details of hospital episodes is necessarily

short;[2]

- the patient or family may often be unaware of the principal diagnosis upon discharge;
- accuracy of recall may well be a function of the respondents' educational level or income class;
- the above problems are compounded in the case of proxy respondents;
- the survey approach is time consuming and very expensive.

In Ontario, surveys to obtain information on hospital use are unnecessary if access to the OHIP discharge records is obtained.

But the OHIP files do not, in isolation, provide a data base with which we can address the questions asked in Chapter 1. It is not enough to know that OHIP family X had a very bad year, incurring sixty days of inpatient hospital care, if we know nothing else about that family. The OHIP files do not contain information on family size and composition (only persons who receive services are registered), family income, or the education and occupation of the head of the family or spouse, to name but a few missing data. In addition the discharge information is recorded by OHIP number. The necessity of maintaining confidentiality inhibits ready access to those records.

The usefulness of the OHIP hospital records for this type of analysis, then, depends on being able to obtain the necessary supplementary information, and on developing a means of merging data sets while maintaining confidentiality. After considering several possibilities we chose to collect the necessary data and identifiers (OHIP numbers) by means of a household survey on a sample of the Ontario population. This allowed us to request the Ontario Ministry of Health to perform an internal linkage, the product of which would be a linked data set, but without the OHIP numbers.

THE OHIP FAMILY - SOCIOECONOMIC AND DEMOGRAPHIC DATA

The household survey design, the process of drawing up an appropriate

2 It is reasonable to suppose that hospitalization and surgical experiences are reported with greater accuracy than visits to physicians or episodes of illness involving no contact with the delivery system, but the underreporting problem still exists (Fuchs 1978).

population sample to whom the survey could be directed, the survey instrument itself, and the various attributes of the completed survey are reported and discussed extensively in Manga (1978). Therefore, only the essentials will be provided here.

The survey included questions on family size, age and sex composition, marital status, education, occupation, family income, residence, distance and travel time to physician and hospital, and indirect costs involved in receiving medical or hospital care (loss of earnings, travel costs, drugs, and so on). Several 'opinion' queries about the availability, quality, and cost of hospital and medical care in Ontario were also contained in the survey. The data were collected for a probability sample of the population of all 'OHIP families' in Ontario. The set of eligible 'OHIP families' was defined to include only those families with a single OHIP number during the twelve months from 1 April 1974 to 31 March 1975. The probability sample was based on a stratified multi-stage cluster sample design. The population of OHIP families was stratified into subpopulations, by family income and by availability of physicians and hospital resources.[3] Within each, probability subsamples of OHIP families were independently taken.

The many stages were necessary because there was no master list of OHIP families. The first stage consisted of obtaining a probability sample of provincial enumeration areas. This was followed by a sampling of addresses within each sampled enumeration area (stage two), a sampling, at specific rates, of households at sampled addresses (stage three), and a sampling of the OHIP families within chosen households (stage four). A number of selected households and OHIP families were found to be ineligible. Of 1867 selected households, 4.6 per cent were found to be 'dead' addresses and 3.2 per cent of the selected OHIP families were ineligible. The final sample consisted of 1290 families. Since our survey deliberately avoided 'epsem' sampling (in which every element has the same probability of appearing in the sample), responses were weighted to

3 Five income groups were used: $0-4999, $5000-9999, $10,000-14,999, $15,999-19,999, and over $20,000. In fact, 'household income' was used as a proxy, but it is highly correlated with family income. Information on average 1971 incomes for each enumeration area was used to allocate OHIP families to one of these five groups, for sampling purposes only. Details of this and of the stratification by supply factors may be found in Manga (1978, Appendix D).

correct for unequal selection probabilities. The weighting process takes account of the fact that our sample has a disproportionately high representation from high- and low-income enumeration areas. Similarly those families in areas deemed to have an inadequate supply of health care resources were oversampled. The weighting also takes account of biases that could arise from varying non-response rates across income classes.[4]

Once the final sample had been constructed and the survey responses gathered and recorded, the survey data were submitted to the Ontario Ministry of Health's Data Dissemination and Evaluation Branch. There the hospital use experiences of the 1290 selected families over the period 1 April 1974 to 31 March 1975 were drawn from the provincial discharge tapes.[5] The two data fields (OHIP and survey) were then assigned random matching identifiers in place of the OHIP numbers to ensure confidentiality of individual sampled families, and the resulting data sets were released by the Ministry for the use of this project.

MEASURING HOSPITAL BENEFITS

With a linked data set providing information on family income, composition, residence, and so on and on family use of hospitals, what remained was to establish a method of evaluating the incidence of hospital benefits. As Manga (1978, 49) noted in the earlier analysis of medical benefit incidence, 'a common solution ... has been to use the cost of producing ... services ... In the case of hospital services, even reasonable cost estimates are not available, because of complex problems in defining and measuring hospital outputs and the related production processes.'

The acute care hospital's primary function is to enhance the health status of those it treats. But while advances are being made in the development of community and individual health status measures (see Culyer (1978) for a description of the state of the art and remaining problems as they relate to Ontario), there are as yet no unanimously

4 Again, details are available in Manga (1978, Appendix D).

5 In fact, all separations in that period formed the hospital use data. Thus the records did not include patients who were admitted during the study period but discharged after 31 March 1975. By the same token, however, a patient who was admitted before 1 April 1974 and discharged before 31 March 1975 and who was a member of one of the chosen OHIP families would have been in the analysis.

accepted health status measures that could be applied to this particular analysis. Furthermore, even with a satisfactory method of measuring health status, the problem remains of differentiating changes in health status resulting from hospital care from natural changes brought about, for example, by the self-limiting nature of an illness.

The difficulties inherent in using the health status concept for hospital output measurement have led to the adoption of a number of proxies - services, discharges, days of care, and expenditures. Each has its attendant problems (Berki 1972), some worse than others. For example, using services poses a formidable aggregation problem, and services are, in any event, inputs into a production process. Discharges or separations embody well-known problems when used as output measures. Perhaps the most severe is that there is wide variation in the lengths of stay for different conditions requiring hospitalization. Yet each discharge is explicitly assigned an equivalent weight of one if total cases is the adopted output measure. The use of 'number of days' involves similar, but perhaps less severe, problems. As Berki (1972, 34) points out, if a treated case requires 'admission-specific,' 'stay-specific,' and 'diagnosis-specific' services, the use of the unweighted separation may accurately capture 'admission-specific' services, but neither of the other two. 'Stay-specific' will depend on length of stay, and 'diagnosis-specific' on the conditions requiring treatment. Unweighted days is also accurate for one out of the three categories - 'stay-specific' services, but fails on 'admission-specific' and 'diagnosis-specific' criteria.

Using hospital costs (expenditures) as an output proxy entails the assignment of explicit weights to a utilization measure. Since expenditures are equal to services × cost per service, or to days × per diem, or to cases × cost per case, with the right choice of weights one can attempt to capture at least two of Berki's three service categories. While expenditures embody a mix of the problems associated with services and discharges or days' stay, this measure starts to adjust for the problems associated with straight unweighted aggregation. In addition, expenditures as benefit measures are consistent with our interest (noted in Chapter 1) in assessing the distribution of ex post risk-bearing taken on by the hospital insurance program. One of the main goals of this study, then, is to develop and compare various weighting mechanisms. Our analysis thus attempts to distribute hospital benefits, defined in terms of

alternative hospital cost estimates, across income classes.

For descriptive purposes at this juncture, the patient day is employed as the use side of expenditures, although similar issues are relevant in using the treated case, separation, or discharge. Both patient days and separations are used in the empirical analysis in Chapter 4. The most common 'cost' measure building on the patient day is the 'per diem,' defined as total operating expense divided by total days' stay. One might employ a province-wide estimate or hospital-specific figures. The use of a provincial per diem, however, assigns equivalent weights to each day of care in the province. Not only does it not differentiate different diagnoses, but it also assumes hospital-invariant treatment costs for a given diagnosis. In short, by adopting the provincial average per diem we learn no more about output distribution than by using unweighted patient days. Our expenditure incidence results reported in the following chapter include estimates based on the provincial per diem as 'benchmark' figures against which to view incremental degrees of sophistication in weighting schemes.

In the results reported in Chapter 4, the hospital-specific per diem forms the weighting for our second measure of expenditure incidence. Employing hospital-specific per diems incorporates the potential for inter-hospital average cost differentials, thereby providing some improvement over the province-wide measure, but it does not address the reasons for those cost differentials. Since the measure of use with which it is being combined very specifically refers only to inpatient care, and since at least some share of those cost differences between hospitals may be due to different mixes of non-inpatient care activities, problems remain. Two patients entering, respectively, a small community hospital and an urban teaching hospital for identical conditions requiring identical lengths of stay would be recorded as having incurred quite different hospital costs if gross hospital per diems were used to weight their stays. Hospital activities commonly include some subset of outpatient care, emergency care, research, and education, and the mix of those activities varies considerably across Ontario hospitals. For example, although most hospitals record no expenses associated with education, for the teaching hospitals the share of total operating costs allocated to education may be as high as 8 per cent (basis 1974).

The deficiencies in the per diem result from the incompatibility of its

numerator and denominator, the latter specifically referring to inpatient care, the former embodying an amalgam of activities. Compatibility requires that an effort be made to isolate the cost of providing inpatient care. Our third and fourth sets of expenditure incidence results are based on such cost estimates - inpatient hospital-specific cost, per case and per day. The method for allocating costs across and between intra-hospital activities is described in detail elsewhere (Barer and Evans 1980). Very briefly, the procedure entailed the extraction of all non-inpatient cost items from total operating expenditures, using data provided by hospitals for Statistics Canada's Annual Return of Hospitals (HS-1 and HS-2). Non-inpatient items included education, special research programs, ambulance, outpatient department, emergency department, and estimated outpatient shares of predominantly inpatient activities such as EEG, ECG, laboratory, and radiology. A pro rata share of administration expenses was also eliminated. For the 182 public acute care Ontario hospitals on which this allocation was carried out, the average estimated inpatient share of total operating expenses was just over 81 per cent.[6] The hospital-specific values ranged from just under 50 per cent for Princess Margaret Hospital (the Toronto cancer hospital, which recorded a large expense component for research), to 100 per cent for a small rural hospital.

Using inpatient operating expenses estimated in this manner as a numerator, we computed weights for our third and fourth expenditure incidence measures. These weights are denoted henceforth as 'inpatient hospital-specific per diem' (IPDHS) and 'inpatient hospital-specific cost per case' (IPCHS).

The final set of benefit incidence measures used in the following chapters attempts to incorporate the diagnostic variability in inpatient hospital costs. Just as it was unrealistic to assume that two patients admitted to different hospitals for like conditions would incur identical costs, so it is equally unlikely that two randomly selected patients admitted

6 While there were over two hundred acute care hospitals in Ontario in 1974, the 182 hospitals formed the data base for computation of the fifth and sixth expenditure weights. Those weights were developed elsewhere (Barer 1981b) in a six-year analysis of Ontario's hospitals that required a common set of hospitals in operation as acute care facilities over the entire period. Only 182 hospitals met those criteria. All patients in our survey sample who received hospital treatment during the 1974-5 period of analysis were treated in one of those hospitals.

to the same hospital will incur the same costs on a daily or episodic basis. A day spent in the operating room undergoing a triple bypass operation is recorded as a hospital day just like a second post-tonsillectomy recovery day. But there the similarity ends, and none of the previous refinements - from provincial per diem, to hospital per diem, to inpatient hospital-specific per diem - captures this dimension of heterogeneity.

In a study done for the Ontario Economic Council (Barer 1981b), a hospital cost model was employed as a simulator to generate diagnosis-specific marginal hospital costs. In that particular application, marginal costs were aggregated over six years and 182 hospitals to generate a single diagnostic vector containing marginal costs in 1974 dollars for 237 Ontario Broad Code categories.[7] It is those cost estimates, based on the same six years, but remaining hospital-specific, that we apply as weights to generate our fifth and sixth benefit measures.

The marginal cost estimates for 1974 were generated in three stages.[8] First, a behavioural model of inpatient costs in Ontario hospitals was articulated. Briefly, it suggested that variations in inpatient costs per case and per day were functions of such exogenous factors as hospital case flow (in fact the inverse of cases per bed year), occupancy rate, extent of educational activity, average wage level, specialization, and patient mix - diagnostic and by age and sex. Second, a time-series/cross-section estimation spanning the period 1969-74 and incorporating 182 hospitals was undertaken to generate maximum likelihood parameter estimates that represented the impact of each exogenous factor in explaining variance across hospitals in inpatient cost per case and per day.[9] The third stage in the generation of the marginal costs used those parameter estimates as the basis for a simple simulation model.

7 In fact, the Ontario Broad Code comprises 260 categories. Owing to the small number of cases in twenty-three categories and the fact that the method used by Barer (1981b) to develop the case mix variable (CMPXC1) for the cost analysis becomes invalid for small numbers, those twenty-three categories were aggregated into the alternative broad code most applicable by aetiology and use of resources.

8 The following is a very sketchy description. The analytic details may be found in Barer (1981b).

9 The estimation method adjusted for auto-correlation in line with Barer's (1981a) results, showing that a pooled time-series/cross-section ordinary least squares technique generates inefficient estimates, because of a significant degree of auto-correlation in residual errors of estimation over time.

By multiplying the estimated value of inpatient cost per day for any hospital by the number of days of care it provided, one can obtain an estimate of total inpatient cost for any given year. The simulation then addresses the following question: What is the impact on total inpatient costs of a decrease in the number of treated cases of a particular diagnosis for that hospital? By recomputing all variables dependent on case mix and case total (such as occupancy rate, case mix complexity, and case flow rate) one can calculate an adjusted average cost estimate. This recomputation builds in an assumption that all cases 'eliminated' would have incurred the average length of stay for that diagnosis in that hospital. The new estimate of average inpatient cost per day, when multiplied by an adjusted total days' stay [actual days - ('eliminated' cases × average length of stay (ALS) for those cases)] yields an adjusted total expense figure. The difference in the two total inpatient cost estimates, when divided by the adjustment to total days, provides an estimated marginal cost of providing a day of inpatient care for cases of that diagnostic category in that hospital in that year.

This procedure was repeated in Barer's (1981b) analysis for each hospital and each of the 237 diagnostic categories for each of the six years, and a comparable process was followed for computing marginal inpatient costs per case. For the purposes of our study, we needed hospital- and diagnosis-specific marginal costs, per case and per day. Accordingly, for each hospital, we aggregated the per diem and per case cost estimates for the six years, using each hospital's intertemporal distribution of days (or cases) as weights. Finally the marginal costs were 'inflated' so as to be expressed in 1974 dollars, using a hospital sector Paasche index described in Barer and Evans (1980) and applied to Ontario by Barer (1981b). These hospital- and diagnosis-specific marginal costs were then linked with the hospital, diagnosis, and days' stay information from the discharge records of families in our household survey. Thus, the marginal costs became the fifth and sixth utilization weights for our analysis of the distribution of hospital benefits. On average, marginal costs turned out to be approximately 75 per cent of average costs (Barer 1981b).

One might argue that average rather than marginal costs would be the more appropriate weights for estimating benefit incidence, since all hospital costs must be allocated somewhere. However, if one thinks of a hospital's fixed costs (the costs of putting in place and staffing a certain number of

beds and providing technical back-up) as satisfying the option demand of the entire community,[10] then those fixed costs (benefits) accrue equally to each member of that community,[11] and any consideration of benefit incidence can quite legitimately focus on the distribution of marginal costs.

SUMMARY

In the absence of prices determined by a market, the estimation of hospital benefit distribution requires not only information on the distribution of hospital use but also a means of inferring a benefit structure from patterns of use. In the context of this research, that implies two major problems. First, we need a method of linking socioeconomic family descriptors to the families' use of hospitals; and secondly, we need a method of computing the benefits of that use.

This chapter has described the data base from which family socio-economic and demographic data and hospital use records were taken and the method of linking them. In addition, considerable attention has been devoted to methods of moving from patterns of use to distribution of benefits. A number of increasingly complex methods of weighting discharges and days' stay with 'pseudo-prices' were described. The intent was to incorporate the fact that costs of hospitalization (and by implication benefits in the sense of private costs avoided) vary across hospitals and diagnoses. We developed six different weighting methods, using as weights provincial per diems, hospital-specific per diems, hospital-specific inpatient per diems, hospital-specific inpatient costs per case, and hospital- and diagnosis-specific per diems and costs per case.

10 Option demand refers to a conditional demand by consumers for available capacity. Although a given population group may not require hospital services in a given period, they nevertheless value the existence of that capacity and derive utility from knowing that, should the need arise for them to exercise effective demand, the appropriate supply-side capacity will be in existence (Weisbrod 1964).

11 This argument abstracts, of course, from inequities of access, which may mean that urban dwellers' option demand, for example, is more fully satisfied than that of rural residents. In other words, option demand is somewhat inappropriately treated as a public good. In addition, at any time one could argue that an individual's option demand is directly related to his probability of requiring the use of the resource. This probability would presumably be related to age, past health history, and so on. But over a lifetime, each person will on average exercise the same 'cumulative' option demand.

4
THE INCIDENCE OF HOSPITAL BENEFITS IN ONTARIO

The potential biases introduced by the use of unadjusted average cost measures to impute hospital benefits from data on use were discussed in the previous chapter. There it was noted that applying province-wide per diems or per case cost measures amounted to equal weighting of each day or case, respectively, and that while using costs to estimate benefits was legitimate under certain assumptions and as a means of answering certain questions, it was necessary to refine the cost estimates. Finally, six different cost measures were described. In this chapter we present and discuss the benefit incidence results based on those six measures.

The results presented in the following pages are based on a straightforward cross-tabulation analysis. Such analyses do not imply anything about the causal structure or direction among the relationships examined. This topic is discussed in Chapter 5. While the causal analysis in that chapter focuses exclusively on the OHIP family, here we consider both family- and person-specific results. This allows some assessment of the role of variables other than those within each cross-classification. For example, variance in hospital expenditures across income classes may result from different average family sizes within different income classes. Reporting per person results by income class may serve to clarify the influence of family size.

All results appearing in the tables of this chapter are based on a sample that was weighted to adjust for the fact that our procedure oversampled families in high- and low-income enumeration areas and in areas that were 'under-supplied' with medical and hospital resources. The rationale for the weighting was considered briefly in the previous chapter, and the process is described in detail elsewhere (Manga 1978). Since the weighting was done on a family basis, the results reported on a per person basis were computed by weighting each person by his or her family

weight factor. This generates some obvious, but small, discrepancies in comparing per family and per person results. For example, the cost estimates for single-person families would be identical, whether reported on a per family or per person basis, if the results were based on an unweighted sample. Similarly, the 'married couple' results on a per family basis would be twice the corresponding per person figures on an unweighted basis. The weighting leads to small differences in the two sets of results.

While the results from this analysis give an indication of the pattern of hospital use and benefits across income classes and other variables describing the population, our interests extend beyond that purely descriptive information. In particular, the principal questions posed at the outset of this study were whether there were significant variations in hospital benefits between different segments of the population and, if so, why. An analysis of variance in the cross-tabulation results indicates the extent and significance of any such variations. An F-statistic is used to test the null hypothesis that the mean benefits or use across socioeconomic groups are identical.[1] F-statistics are reported for all comparisons in this chapter.

HOSPITAL BENEFIT INCIDENCE IN ONTARIO

For each person from our sample of OHIP families who was discharged from a hospital in Ontario from 1 April 1974 to 31 March 1975, information was available identifying the hospital to which the patient was admitted, the patient's age, sex, length of hospital stay, and the principal diagnosis on

1 The F-statistic compares the sample variance between means (its numerator) with the variances around each of the mean values (the denominator). If the former is small in relation to the random fluctuations within the socioeconomic groups, we cannot reject the null hypothesis. But if the variation between means is large in relation to the variances within sub-samples, we can be fairly confident that the differences in sample means are the result of different underlying population means (i.e. the null hypothesis ought to be rejected), rather than of random unexplained variance. While analysis of variance assumes normally distributed homoscedastic subsamples, it was pointed out elsewhere that the technique 'is generally considered to be sufficiently "robust"... to provide a useful approximation even for wide departures from both normality and homoscedasticity' (article by Suits referenced in Manga 1978, 60).

discharge. Of the 1290 families in the sample, 377 had one or more discharges. Among the persons in the sample during the year, there were a total of 624 weighted discharges, requiring 3854 weighted days of hospitalization for an average length of stay of 6.2 days. The weighted number of families with at least one discharge was 360.

The Ontario-wide average length of stay of 9.6 days for patients discharged from acute care hospitals in 1974 (Ontario n.d. 1976, 20) was 55 per cent higher than the average for our sample. This large discrepancy, which raises some doubt about the degree to which our sample is representative of the Ontario population,[2] seems to result from two interacting factors, one quantifiable, one not. First, the hospitalization records for our sample include only discharges from acute care units of public hospitals, although in 1974 there were ninety-six chronic units and eleven rehabilitation units in the 194 acute care hospitals in Ontario. The average length of stay for the active units was 8.8 days, which is a still considerable 42 per cent higher than the average for our sample.

Secondly, the survey procedure was such that families with persons hospitalized at the time of the household survey would have been more likely to be left out of the final operational sample than families in which no one was hospitalized. In particular, a single person targeted in the original sampling frame but in hospital at the time of the field contacts would have been missed. The longer the hospital stay, the more likely that such a person would not have been included. While this problem is most severe for single-person families, it is a possible explanation for the relatively low average lengths of stay in all family classifications. If one member of a married couple was in hospital, the other was likely to be visiting or working and might therefore have been excluded from the sample. A single parent who was hospitalized would almost certainly have been missed, and if one or more of the children was hospitalized the family would again probably have been missed. We do not know to what extent this sampling 'bias' explains the difference in average length of stay

2 In fact because of the unique definition of the OHIP family, and the lack of the requisite population data, we are unable to determine how representative our sample is of the 1974 Ontario population as regards income class, family composition, and age of family head. There is no 'master list' or registration or census for OHIP families, and even with such a list, information on income-class distribution would still not be available.

between 6.2 days and 8.8, but it is safe to say that any family with at least one member in hospital over the entire time of the field survey (including recalls) had a better than average chance of being excluded from the final sample. The chances of missing the original contact plus the follow-ups increased, of course, with the length of the hospital stay. In fact, our sample did not include any patients with lengths of stay of over thirty-five days, despite the fact that many active treatment units house a few chronic care patients who incur very long stays.

A quick check of the distribution of separations for our sample showed that there was one or more separations for only 148 out of the 237 adjusted Ontario Broad Code (OBC) categories (Barer 1981b). Simple correlations of our sample's diagnostic distribution of separations and average length of stay with those of the Ontario population in 1974 were 0.75 and -0.01 respectively. When the OBCs categories that contained no one from our sample were eliminated, the correlations for the vectors based on the remaining adjusted OBCs were 0.72 and 0.36 respectively. Unfortunately that is the only information that is in any sense 'matchable' with the general population, and it does not tell us anything about the representativeness of our sample in terms of the socioeconomic factors of importance to this analysis. What it does seem to suggest is that, whereas the mix of diagnoses is relatively representative, the lengths of stay within diagnoses are not, a finding not inconsistent with the reasons noted above for the sample's shorter length of stay than the provincial average.

Although the different average lengths of stay for the total population and for our sample cannot, therefore, be precisely reconciled, we can infer something about the likely effects of the difference on our cross-tabulation results. The largest proportions of single-person families were in the lowest income classes. The large OHIP families (married couples with three or more children) are found predominately in the upper-income classes in our sample. Given our second point above, this would suggest that the hospitalization pattern for our sample will be biased downward in relation to the population for the lowest income classes. The bias is likely to shrink with increased income.

Similarly, we might expect downward-biased hospitalization results for single-person families and a similar but less severe relative bias for married couples with no children and for single parents. As regards the age of the head of the household, since a relatively large proportion of the

65-and-over group were also single persons, the relative hospitalization of this group will again be understated by our sample.[3]

The cross-classification results are reported for two measures of use (days' stay and separations) and six expenditure (benefit) measures. The expenditure measures are denoted as follows:

- days weighted by the Ontario-wide 1974 hospital per diem - CPDOW;
- days weighted by hospital-specific per diems - CPDHS;
- days weighted by inpatient hospital-specific per diems - IPDHS;
- cases weighted by inpatient hospital-specific average costs per case - IPCHS;
- cases weighted by inpatient hospital-specific and diagnosis-specific marginal costs per case - MCC; and
- days weighted by inpatient hospital-specific and diagnosis-specific marginal costs per day - MCD.

Of course the distribution of days' stay and CPDOW will be identical, the latter being the multiple of the former where the multiplicative factor is the Ontario average per diem for 1974 - $105.91 (Ontario n.d. 1976, 110). The tables present the distribution of these eight variables by family income class, age of the head of the household, and family composition.

In Table 11 the utilization results are presented on a family basis for these three sets of socioeconomic classifications. On average for the entire sample, one hospitalization experience occurred for every two families, and each family required hospitalization for three days. Column 2 indicates no significant income-related differences in the number of separations per family. Families in the lowest income class did require more hospital days than their better-off counterparts (6.07 compared to the average of 2.99, column 1). The highest income class shows the fewest average hospital days (2.27), while the $4000-7999 income class had an atypically low

3 In Chapter 2 we noted that the Statistics Canada (1977) study also underestimated the actual hospitalization of the Canadian population. We suspect that among the factors responsible for this discrepancy is, again, the non-sampling of the hospitalized population. Studies such as Beck and Horne (1980) and Horne (1978) avoid this problem because of larger sample sizes and because they did not use surveys.

TABLE 11
Average hospitalized days and separations per family

	Total sample: Days' stay (1)	Total sample: Separations (2)	Percentage of families with one or more separations (3)	Families with one or more separations: Days' stay (4)	Families with one or more separations: Separations (5)
Overall average	2.99	0.48	29.2	10.69	1.73
Income (INC)					
$ 0- 3,999	6.07	0.46	27.0	20.89	1.58
4,000- 7,999	2.34	0.24	24.4	13.18	1.37
8,000-13,999	3.16	0.64	29.0	11.72	2.36
14,000-19,999	2.40	0.48	30.0	7.84	1.56
20,000 +	2.27	0.41	26.6	7.22	1.30
Unknown	2.95	0.52	34.0	6.87	1.20
F-statistic	3.58*	1.35		5.79*	1.51
Age of head of household (HEADAGE)					
18-34	2.51	0.52	30.4	7.91	1.64
35-49	2.99	0.65	31.7	9.36	2.04
50-64	3.31	0.36	27.5	13.30	1.44
65 +	3.50	0.24	25.0	22.16	1.53
F-statistic	0.65	2.86**		8.47*	0.79
Family composition (FAM)					
Married couple	2.54	0.25	22.3	12.83	1.28
Married couple with 1 or 2 children	3.70	0.60	35.6	10.90	1.76
Married couple with 3 or more children	4.49	1.07	39.6	12.93	3.08
Single person	1.89	0.25	16.2	8.49	1.13
Single parent with children	2.39	0.53	36.0	6.15	1.37
F-statistic	3.15**	7.71*		1.83	4.10*

* Significant at 0.01 level
** Significant at 0.05 level
Significance indications based on two-tailed tests

average separation rate (one-half of the sample average of 0.48).[4]

Column 3 provides a transition from the whole sample to the sub-sample of user families only. It shows that the proportion of families with at least one hospitalization did not vary appreciably across income classes. The families in the $4000-7999 class were least likely to have experienced hospitalization, a fact that partly explains why that class had the lowest separation rate based on the whole sample.

When we consider only the families with at least one separation, the average number of hospital days was 10.69 and the average number of separations was 1.73. As for the entire sample, again there was no significant income-related trend for separations (column 5). However, a significant inverse relationship is noted between the average number of hospital days and income class (column 4). In fact, families in the lowest group spent, on average, almost three times as many days in hospital as families earning $20,000 or more.

Turning to the results by age of the household head, we observe no significant differences in days' stay when the total sample is considered (column 1), but the differences in average separations are significant (at a 95 per cent confidence level; see column 2). In particular, younger families have more frequent separations, a result of a large proportion of those families having at least one separation (column 3). This trend is a function of family size, since family heads in the 18-49 age bracket have larger families than the more elderly heads. Column 3 also reflects the fact that our study is limited to acute care hospitals. The aged are often institutionalized for treatment or convalescence in long-term care hospitals, nursing homes, psychiatric hospitals, and so on, or are cared for at day hospitals or through home care programs. In fact, the significant differences in separations (column 2) disappear when the more restricted

4 The following table contains the degrees of freedom for determining critical levels for the F-tests:

	All families	Families with one or more discharges	All persons	Persons from families with one or more discharges
INC	5/1284	5/371	5/3533	5/1099
HEADAGE	3/1286	3/373	3/3535	3/1101
FAM	4/1285	4/372	4/3534	4/1100

The 'Unknown' income category comprised 44 survey respondents (out of the total of 1290) who declined to provide this information.

sample is viewed (column 5). For the families with at least one separation, a strongly significant difference is noted in days' stay, which increases with the age of the family head (column 4). Evidently, for the 'separated' sample, the much longer stays (Table 12) for members of the older families overwhelm the trend in column 3 and produce threefold differences in average days' stay per family for persons 65 and over as compared to the 18-34 HEADAGE group.

The family composition results indicate significant differences in separations as one moves across different family compositions (columns 2 and 5). Of course one would be quite surprised not to find that larger families have more hospitalization per family. As for days' stay, the significant difference evident for the whole sample (column 1) disappears when only the 'separated' sample is considered (column 4), because there are more families with no hospital separations among the smaller families (column 3). Thus we see the greatest increase in days' stay, when comparing the total sample with the more restricted subsample, for single-person families and the second largest increase for married couples.

As noted earlier in this chapter, the average length of stay (ALS) per separation for patients in our sample is approximately 6.2 days. In column 3 of Table 12 this average length of stay is disaggregated by our socioeconomic groups. ALS generally declines with increases in income. In particular, ALS for families in the lowest income class is over twice the overall average, while there is little to separate the upper three income classes. The ALS increases with age of the family head and is highest for married couples. The lowest ALS by family composition is found in the largest families, undoubtedly reflecting short stays by children. The ALS figures in Table 12 explain the differences in patterns observed in columns 4 and 5 of Table 11. In fact, ALS = column 4/ column 5.

The difference between columns 1 and 2 of Table 12 and the results in Table 11 is simply one of denominators. That is, in Table 11 the denominator for all the figures is the number of families, whereas in Table 12 it is number of individuals. For large families, this implies a larger denominator in the Table 12 figures than in those of Table 11. This is illustrated most dramatically by the 12.93 days for the married couples with three or more children in the 'separated' sample (Table 11), which falls to 2.48 days per person in Table 12.

It is evident from columns 1 and 2 of Table 12 that on a per person

TABLE 12
Average hospitalized days per person and average length of stay

	Days' stay per person		
	Total sample (1)	Persons from families with one or more separations (2)	Average length of stay per separation (3)
Overall average	1.09	3.49	6.18
Family income (INC)			
$ 0- 3,999	4.05	11.76	13.22
4,000- 7,999	1.14	5.20	9.62
8,000-13,999	1.19	4.14	4.96
14,000-19,999	0.70	2.18	5.04
20,000+	0.68	1.94	5.55
Unknown	1.14	2.27	5.71
F-statistic	18.27*	21.68*	
Age of family head (HEADAGE)			
18-34	0.94	2.47	4.82
35-49	0.83	2.75	4.59
50-64	1.38	5.07	9.23
65 +	2.21	10.84	14.5
F-statistic	8.00*	23.73*	
Family composition (FAM)			
Married couple	1.31	6.55	10.20
Married couple with 1 or 2 children	0.97	3.19	6.18
Married couple with 3 or more children	0.96	2.48	4.19
Single person	1.89	8.42	7.54
Single parent with children	0.73	1.61	4.50
F-statistic	3.45*	16.13*	

* Significant at 0.01 level
Significance indications based on two-tailed tests

basis there are more marked differences in days' stay between socio-economic groups. To be particularly noted are the HEADAGE results for the total sample and the FAM results for the hospitalized subsample. On a per family basis for the whole sample (Table 11, column 1), we were un-

able to reject the hypothesis that mean days' stay was identical when the age of the head of the household was the variable of comparison. The comparable per person results, however, indicate significantly more hospitalized days for persons of aged families (particularly with heads of 65 years and over). This is simply a reflection of the generally inverse relationship between age of head of household and family size. The significant differences for the hospitalized subsample across FAM groups, which did not appear in the per family results (Table 11, column 4), are also a function of disparate family sizes. In particular, single-person families had the second fewest days' stay on a per family basis. But since that days' stay figure remains essentially unchanged when we consider the results on a per person basis, the single-person family ends up with the highest number of days' stay per person among families with separations (Table 12, column 2). Probably this is partly due to the lack of someone to care for the patient at home. Also, many of the single-person OHIP families are more than 65 years of age.

These basic cross-classification results reveal that the poor are greater users of acute care hospital facilities than persons or families in all other income groups. This is partly a result of the age mix of that income category and, in turn, average lengths of stay for the elderly. In addition, family composition will have some role to play because a certain number of elderly persons who incur long hospital stays will be widowed or separated - single-person families. These figures alone, however, do not provide a basis from which to determine the relative <u>independent</u> roles of age, family composition, and income in explaining the significant utilization differences across income classes. That must be left to the following chapter.

Tables 11 and 12 have reported hospital use by our total sample and by the subsample of families with one or more discharges, on a per family basis (Table 11) and per person basis (Table 12), as illustrated in the following matrix:

	Total sample	Subsample of families with separations
Per family	Table 11 Table 13	Table 11 Table 14
Per person	Table 12 Table 15	Table 12 Table 16

The next four tables present the benefit incidence results based on our six different methods of weighting use with cost estimates. As illustrated above, there is one such table for each of the 'cells' in the matrix. Thus, Table 14, for example, presents expenditure estimates on a per family basis for the families with one more discharges.

Table 13 illustrates clearly the effect of applying alternative weightings to the hospital use of our whole sample. As noted earlier, the average costs in the CPDOW column result from multiplying the days of care per family for the entire sample by $105.91, the 1974 Ontario-wide per diem for acute care hospitals (Ontario n.d. 1976, 102). If the entire population of Ontario had formed our 'sample,' the per family cost estimates based on CPDOW and CPDHS (hospital-specific per diems) would have been identical. The fact that the latter estimate ($276.72) is lower than the former ($317.04) in Table 13 means that, on average, the hospitals where our sample of families was treated had lower costs than average. Since the sample of 1290 families is relatively small, and since there were fewer than 700 separations for our entire sample, it is not particularly surprising that the two figures are not closer.

The relationship between the estimates based on IPDHS (hospital-specific inpatient per diems) (column 4) and CPDHS (column 2) reflects the results of using the inpatient cost estimation method described briefly in the previous chapter (and in more detail in Barer and Evans 1980). For the set of Ontario hospitals to which that method was originally applied (Barer 1981b), estimated inpatient costs as a share of total operating expenses clustered heavily in the 0.80 to 0.85 range. The ratio of the average IPDHS estimate to its CPDHS counterpart for the hospitals used by our sample is in the middle of that range, suggesting that whereas the hospitals used by the sample may have had relatively low costs, their mix of inpatient/non-inpatient activities conformed closely to the provincial norm.

The implications of using number of separations weighted by costs per separation as an estimate of hospital benefits or expenditures is illustrated vividly by the overall average figures in Table 13 (columns 3 and 5) and is reinforced time and again throughout this and the following three tables. The overall estimate based on IPCHS is nearly 50 per cent higher than the one based on IPDHS. Which is 'more correct'? For our particular sample there are two problems in using per case cost measures as weights in

Cost of inpatient care per family, total sample

	Expenditure weight					
	CPDOW (1)	CPDHS (2)	IPCHS (3)	IPDHS (4)	MCC (5)	MCD (6)
Overall average	317.04	276.72	335.87	228.53	148.68	118.79
Income (INC)						
$ 0- 3,999	642.83	518.53	359.38	434.59	150.49	216.20
4,000- 7,999	247.70	222.67	185.70	183.83	86.97	95.95
8,000-13,999	334.47	293.35	416.41	241.22	187.58	130.74
14,000-19,999	254.41	238.31	330.66	194.52	142.19	98.69
20,000+	240.78	202.20	293.60	168.70	124.54	87.44
Unknown	312.02	270.18	354.62	224.38	185.42	109.96
F-statistic	3.58*	2.55**	1.01	2.78**	1.15	2.31**
Age of head of household (HEADAGE)						
18-34	266.07	221.65	348.48	184.38	148.43	97.83
35-49	316.67	296.10	453.33	241.56	195.25	124.63
50-64	350.11	301.65	245.98	250.66	120.84	128.77
65 +	370.46	306.83	185.38	254.66	88.97	133.07
F-statistic	0.65	0.74	2.75**	0.74	2.20	0.57
Family composition (FAM)						
Married couple	269.03	236.38	189.17	193.79	87.20	102.02
Married couple with 1 or 2 children	392.01	323.60	409.62	271.48	186.94	142.32
Married couple with 3 or more children	475.90	487.54	753.41	387.85	312.94	203.36
Single person	200.45	159.59	173.34	135.64	72.49	67.41
Single parent with children	252.69	218.17	345.39	180.30	157.83	90.64
F-statistic	3.15**	4.48*	8.11*	4.19*	7.80*	4.14*

* Significant at 0.01 level
** Significant at 0.05 level
Signficance indications based on two-tailed tests

moving from use to benefits. The fact that the average based on IPCHS is larger than that based on IPDHS shows clearly that the patients of this sample who were hospitalized stayed in hospital for shorter lengths of time than the average for the hospitals to which they were admitted.[5] Weighting this sample's separations by per case cost estimates derived from a patient load with much longer average stays yields estimates based on IPCHS that are higher (on average) than those based on IPDHS.[6]

More generally, the per case cost measures allow for variance across hospitals only (IPCHS) or hospitals and diagnoses (MCC). The method used to compute those cost measures could, in theory, have disaggregated costs further by socioeconomic class. As noted earlier, however, it is impossible to link socioeconomic information for all OHIP families to the province's hospitalization data. Therefore, for example, MCC implicitly assumes constant average length of stay for any given diagnosis across any set of socioeconomic classes. But as Table 12 illustrated, there are wide variations in length of stay between these classes.

Our final two weights, MCC and MCD, attempt to incorporate differences between socioeconomic classes in diagnostic mix. They are not only inpatient- and hospital-specific, but also diagnosis-specific cost estimates. In addition, as noted in the previous chapter, they are marginal rather than average costs. The relationships between IPCHS and MCC and between IPDHS and MCD in Table 13 show that the marginal inpatient costs are approximately 45 to 55 per cent of average inpatient costs. While the overall average estimates still differ according to whether the basis for estimation was the per case or per day marginal costs, the divergence is narrower than the estimates based on the average inpatient costs. This suggests that in the two estimates (based on IPCHS and IPDHS), some of the divergence generated by length of stay did indeed

5 Indeed, this is reflected in the sample's overall ALS of 6.2 days, itself around 30 per cent shorter than the provincial average for Ontario's acute care hospitals in 1974.

6 It does not follow that the estimates based on IPCHS are biased upwards. It is generally acknowledged that the early days of a hospital stay are, on average, the relatively resource-intensive, expensive days and that many long stays involve predominantly low cost 'maintenance' care. The hospital-specific estimates of per case cost (IPCHS) will of course embody all those long stays. For our relatively short-stay sample, then, IPDHS may underestimate the inpatient per diem cost of their treatment. Or it may not, for factors other than length of stay alone come into play.

result from the lack of diagnostic standardization in those measures.[7]

Let us turn now to the socioeconomic disaggregations in Table 13; in the section on income, the CPDOW column is simply a multiple of column 1 of Table 11. While the relative positions of the income classes are identical in the CPDOW and CPDHS columns, a comparison of those two columns does reveal some interesting information. The difference in the F-statistics for the two columns shows that we cannot be as confident in rejecting the null hypothesis when the estimates based on CPDHS are used. When one compares columns 1 and 2, two cells stand out. While each income class in our sample was treated, on average, in relatively low-cost hospitals, this was particularly pronounced for the lowest income group and least evident for the $14,000-19,999 income class. For the lowest income group, the column 2 figure is 19.3 per cent lower than the figure in column 1, whereas the corresponding reduction for the $14,000-19,999 income class is only 6.3 per cent. It would seem, then, that although the poor are relatively high users of hospital resources, they are also relatively frequent occupants of beds in lower-cost acute-care institutions. The estimates based on IPDHS provide no significant new information at this stage. It appears that the average ratio of inpatient to total operating costs of around 0.83 noted earlier also characterizes the subsets of hospitals treating each income class. Similarly, only minor cost differences based on diagnosis materialize in a comparison of columns 4 and 6. The lowest income group appears to be treated for slightly less costly diagnoses than the total sample, whereas the reverse is true for the $8,000-13,999 class. But the differences are small. For the whole sample, the ratio of benefits based on MCD to those based on IPDHS is 0.52. The corresponding ratios for the lowest income and the $8,000-13,999 classes,

7 This marginal/average differential comprises two components. As noted in Chapter 3, the ratio of marginal cost to average cost for the provincial diagnostic mix was around 75 per cent in 1974. The fact that the ratios of columns 5 to 3 and 6 to 4 in Table 13 are between 45 and 55 per cent shows that the diagnostic mix of our sample was less expensive than the provincial diagnostic mix. In particular, the patients in our sample who were treated in any given hospital were treated for lower-cost diagnoses than the actual mix of patients in that hospital. For a general review of other literature that has compared marginal and average costs, see Lipscomb et al. (1978). Barer and Evans (1980) suggest a reconciliation of discrepancies between ratios based on time-series as opposed to cross-section analyses.

respectively, are 0.50 and 0.54.

Quite different results are generated by the per case estimates. They suggest no significant differences between income classes in hospital benefits, primarily because these weights do not include differences in average length of stay across socioeconomic groups. But the relatively constant ratio of column 5 to 3 does tell us that the differences between columns 1 and 3 are not confounded by any extraordinary diagnostic variation across the income classes. The difference in columns 1 and 3 simply reflects the fact that average lengths of hospitalization were significantly higher than the sample mean for the lowest income groups, and lower for the other groups.

Overall, the INC results in Table 13 suggest that for our entire sample the differences in hospital benefits across income classes are at best (or worst, depending on one's perspective) marginally significant at a 95 per cent confidence level.

Turning to the HEADAGE results, the null hypothesis that hospital benefits per family do not vary with the age of the household head cannot be rejected. The only column in which significant ($p < .05$) differences appear is the one based on IPCHS. That column serves to highlight another problem related to length of stay that occurs when per case estimates are used. As noted earlier, there are significant differences in average length of stay between different family age groups. But the per case measures do not allow that variation to be incorporated in the results in columns 3 and 5 of Table 13. In columns 1, 2, 4, and 6, even when that variance is incorporated, the aged families still do not show significantly higher hospital expenditures (because there are many more families in that age group with no hospitalization, as shown in Table 11, column 3). In other words, when the per case weights are used, the aged families appear to be relatively low beneficiaries of hospital care because they have relatively few separations from acute care facilities and because the IPCHS and MCC measures do not incorporate the fact that those few separations come at the end of lengthy hospital stays.

A comparison of columns 1 and 2 reveals that families with heads 35-49 years of age were generally treated in higher-cost hospitals than other families in our sample. This is shown by the fact that the column 2 estimate for this group is only 6 per cent lower than its column 1 estimate, whereas the reductions for the other groups range from 14 to 17 per cent.

The inpatient cost isolation (column 4) and diagnosis-specific cost estimates (column 6) do not add significant information about distribution. The ratios of column 4 to 2 and column 6 to 4 are almost constant across HEADAGE categories.

The last (FAM) section of this table shows that the probability of erring by rejecting our null hypothesis is smaller than indicated by the estimates based on CPDOW. The differences in benefits (between different kinds of family composition when the entire sample is considered) are more significant than simple differences in use. The differences between columns 1 and 2 are again quite revealing. In particular, the ratio of column 2 to 1 for married couples with more than two children is greater than 1.0, in contrast to ratios ranging from 0.80 to 0.88 for the other compositions. Recall that our sample, on average, was treated in hospitals with less than provincial average costs. Clearly, the largest families in our sample were an exception. Single persons, on average, received treatment in the lowest-cost hospitals used by the families in our sample. As for the relative advantages of the progressively sophisticated estimates based on expenditures, once again little additional ordinal or cardinal information is obtained by moving from column 2 to 4 and then to 6.

Unlike Table 13, Table 14 deals only with the restricted subsample of families with at least one hospital separation. For all the benefit measures based on per day weights (columns 1, 2, 4, and 6), there is less than a 1 per cent chance of erring by rejecting our null hypotheses. Among the families that had at least one hospitalization, it seems clear that there are significant differences in hospital benefits, and that benefits are inversely related to income class. The HEADAGE results show markedly different benefit levels between families of different age structures. This is a product of the variation in length of stay (Table 12), since there were no significant differences in separation rates (Table 11, column 5). Not surprisingly, for families with some hospital experience, the hospitalization benefits increase dramatically with the age of the family head.

On the basis of columns 1, 2, 4, and 6, married couples with no children show greater expenditures than married couples with one or two children. While the latter group shows more separations because of sheer numbers, its members also stay in hospital fewer days once admitted, and this length of stay effect obviously outweighs the separations differential. A similar explanation lies behind the finding that single-person families

TABLE 14
Cost of inpatient care per family, families with one or more separations

	Expenditure weight					
	CPDOW (1)	CPDHS (2)	IPCHS (3)	IPDHS (4)	MCC (5)	MCD (6)
Overall average	1132.45	988.43	1199.73	816.31	531.09	424.31
Income (INC)						
$ 0- 3,999	2211.89	1784.19	1236.58	1495.38	517.83	743.92
4,000- 7,999	1395.81	1254.78	1046.42	1035.93	490.08	540.69
8,000-13,999	1241.01	1088.44	1545.14	894.99	695.99	485.08
14,000-19,999	830.10	777.56	1078.88	634.67	463.92	322.02
20,000+	764.34	641.88	932.03	535.54	395.36	277.58
Unknown	727.15	629.65	826.42	522.92	432.10	256.25
F-statistic	5.80*	4.29*	1.10	4.69*	1.32	3.98*
Age of head of household (HEADAGE)						
18-34	837.63	697.80	1097.07	580.47	467.29	307.98
35-49	991.53	927.13	1419.43	756.36	611.36	390.23
50-64	1408.71	1213.71	1001.79	1008.57	486.23	518.13
65 +	2347.17	1944.05	1187.20	1613.49	563.71	843.09
F-statistic	8.47*	6.88*	0.87	7.36*	0.68	6.38*
Family composition (FAM)						
Married couple	1358.46	1193.61	955.22	978.57	440.32	515.18
Married couple with 1 or 2 children	1154.19	952.77	1191.32	799.34	550.40	419.04
Married couple with 3 or more children	1369.07	1402.56	2167.42	1115.77	900.28	585.03
Single person	899.34	716.04	777.72	608.58	325.23	302.45
Single parent with children	651.07	562.12	889.90	464.53	406.66	233.54
F-statistic	1.83	3.05**	4.57*	2.73**	4.11*	2.74**

* Significant at 0.01 level
** Significant at 0.05 level
Significance indications based on two-tailed tests

incur higher costs than families with children. The larger family group has more separations (Table 11, column 5), but shorter stays (Table 12, column 3). In fact, the families with children could be expected on average to have many short stays resulting from childhood ailments that are often acute but short-lived. In addition, since some nursing care is available at home for those children, some of the recuperative days that a single person might spend in hospital could be comfortably spent at home by children, with little or no difference in the efficacy of care.

Table 15 begins our examination of benefit incidence for the individual rather than the family. Differences between Tables 13 and 15 reflect nothing more than differences in average family size between socioeconomic groups. On a per person basis, significant income class differences result from the extremely high costs incurred by members of OHIP families in the lowest income group. These differences are far more significant than those of Table 13 because the poor families are smaller on average than the families in all other income classes.[8] This, in turn, reflects the disproportionate share of elderly persons (predominantly married couples or single-person families) in that income group. The smaller than average size for elderly families also gives rise to the significant differences between HEADAGE categories (when per diems are the basis for cost estimates), which did not appear in Table 13. On a per person basis, members of families with heads of household over 50 years of age use increasing amounts of hospital care owing to increasing average lengths of stay. Finally, on a per person basis, single-person families are the highest users among our family composition groups.

Table 16 combines the results contained in Tables 12 and 15. Thus, while it continues to report per person results, it does so for the restricted sample of persons from families with at least one separation. The differences in F-statistics and patterns between Tables 15 and 16 are the ones seen in moving from column 1 to column 2 in Table 12, and the discussion of those results is pertinent here as well.

Table 17 presents results based on a further paring of the sample. In moving from Tables 16 to 17, we have eliminated from the denominators

8 In fact, 66 per cent of the OHIP families in the lowest income class were single persons and the average family size was 1.5. For the $14,000+ income class, however, the average family size was over three (Manga 1978).

TABLE 15
Cost of inpatient care per person, total sample

	Expenditure weight					
	CPDOW (1)	CPDHS (2)	IPCHS (3)	IPDHS (4)	MCC (5)	MCD (6)
Overall average	115.33	100.66	122.18	83.13	54.08	43.21
Income (INC)						
$ 0 - 3,999	428.79	345.87	239.72	289.89	100.38	144.21
4,000- 7,999	120.98	108.75	90.69	89.78	42.48	46.87
8,000-13,999	125.98	110.49	156.84	90.85	70.65	49.24
14,000-19,999	74.65	69.92	97.02	57.07	41.72	28.96
20,000+	71.60	60.13	87.31	50.17	37.04	26.00
Unknown	120.49	104.33	136.94	86.65	71.60	42.46
F-statistic	18.27*	14.21*	3.72*	15.17*	4.02*	13.48*
Age of head of household (HEADAGE)						
18-34	100.07	83.37	131.07	69.35	55.83	36.80
35-49	88.36	82.62	126.49	67.40	54.48	34.78
50-64	145.84	125.65	103.71	104.42	50.34	53.64
65 +	234.11	193.90	118.41	160.93	56.22	84.09
F-statistic	8.00*	6.19*	0.40	6.57*	0.09	6.31*
Family composition (FAM)						
Married couple	138.27	121.52	97.05	99.62	44.71	52.43
Married couple with 1 or 2 children	102.36	84.04	103.15	70.67	47.84	36.74
Married couple with 3 or more children	101.89	101.02	153.85	80.88	64.81	42.85
Single person	200.09	159.31	173.03	135.40	72.36	67.29
Single parent with children	77.69	67.08	106.20	55.44	48.53	27.87
F-statistic	3.45*	2.61**	2.17	2.80**	1.44	2.47**

* Significant at 0.01 level
** Significant at 0.05 level
Signficance indications based on two-tailed tests

TABLE 16
Cost of inpatient care per person, families with one or more separations

	Expenditure weight					
	CPDOW (1)	CPDHS (2)	IPCHS (3)	IPDHS (4)	MCC (5)	MCD (6)
Overall average	369.39	322.41	391.33	266.27	173.23	138.40
Income (INC)						
$ 0 - 3,999	1245.92	1005.00	696.54	842.32	291.68	419.03
4,000- 7,999	550.67	495.03	412.83	408.69	193.34	213.31
8,000-13,999	438.76	384.82	546.25	316.43	246.07	171.50
14,000-19,999	230.52	215.93	299.61	176.25	128.83	89.43
20,000+	205.92	172.93	251.10	144.28	106.52	74.78
Unknown	240.54	210.88	276.78	175.13	144.72	85.82
F-statistic	21.68*	17.02*	4.96*	18.20*	5.48*	16.47*
Age of head of household (HEADAGE)						
18-34	261.37	217.73	342.32	181.12	145.81	96.10
35-49	291.61	272.67	417.46	222.45	179.80	114.77
50-64	532.23	458.56	378.49	381.05	183.70	195.76
65 +	1148.11	950.92	580.71	789.23	275.73	412.39
F-statistic	23.73*	19.59*	1.19	20.63*	1.84	19.69*
Family composition (FAM)						
Married couple	693.31	609.29	486.60	499.51	224.20	262.90
Married couple with 1 or 2 children	338.30	277.75	340.91	233.55	158.09	121.42
Married couple with 3 or more children	262.37	260.12	396.16	208.28	166.87	110.35
Single person	892.12	710.29	771.48	603.69	322.62	300.02
Single parent with children	170.06	146.83	232.45	121.34	106.22	61.00
F-statistic	16.13*	12.75*	4.28*	13.72*	3.90*	12.59*

* Significant at 0.01 level (based on two-tailed tests)

all members of families with at least one separation who themselves received no hospital care. The most marked effect of this restriction appears in the FAM results, where strongly significant differences (Table 16) based on the various per diem cost estimates ($p < 0.001$) are reduced to the point where the null hypothesis can be rejected with somewhat less, although still considerable, confidence ($p < 0.05$). This occurs, of course, because the families with children tend also to be the families with the most non-hospitalized persons. Thus the column 6 figures for married couples with three or more children show the most dramatic proportionate increase in moving from Table 16 to Table 17. The two highest FAM figures in Table 16 - married couples and single persons - change the least from the shrinking of the sample (indeed single-person family denominators naturally do not change), while the other figures increase sharply.

Whereas magnitudes of F-statistics also fall in almost all cases throughout the rest of Table 17 (in relation to Table 16), strongly significant differences in hospital benefits remain across income classes and family age categories.

What is interesting about Table 17 is that, when the confounding factor of variation in family size (and thus in the number of potential non-hospitalized persons) is removed, hospitalized persons in the lowest income group tend to incur the highest costs, and as family income increases, incurred costs per hospitalized person fall. Once again the main explanation is found in the HEADAGE figures. As a result of their above-average lengths of stay and their dominant place in the lowest income class, the elderly tend to be an overpowering influence in the benefit distributions by income class and by family composition.

SUMMARY

This chapter has reported use and benefit incidence results for a number of socioeconomic disaggregations. The principal finding was that poor families use significantly more hospital care than families in other income classes, in spite of the fact that the differences between income classes in separations per family are not significant. This greater use by the poor is explained primarily by their relatively high average length of stay, which was more than twice the average for the sample. While there are no doubt a number of reasons for this difference, it is partly due to the

Cost of inpatient care per person, persons with one or more separations

	Expenditure weight					
	CPDOW (1)	CPDHS (2)	IPCHS (3)	IPDHS (4)	MCC (5)	MCD (6)
Overall average	866.52	756.32	917.99	624.61	406.37	324.67
Income (INC)						
$ 0 - 3,999	1760.59	1420.15	984.27	1190.27	412.17	592.13
4,000- 7,999	1071.74	963.45	803.47	795.41	376.29	415.16
8,000-13,999	874.10	766.64	1088.24	630.39	490.22	341.66
14,000-19,999	666.57	624.39	866.35	509.64	372.53	258.58
20,000+	597.99	502.18	729.18	418.98	309.31	217.16
Unknown	633.88	548.88	720.41	455.84	376.67	223.38
F-statistic (degrees of freedom: 5465)	9.30*	6.65*	1.15	7.29*	1.44	6.32*
Age of head of household (HEADAGE)						
18-34	582.17	484.98	762.49	403.44	324.78	214.06
35-49	785.82	734.79	1124.95	599.45	484.53	309.27
50-64	1189.05	1024.46	845.58	851.30	410.41	437.34
65 +	1809.77	1498.95	915.38	1244.07	434.64	650.06
F-statistic (degrees of freedom: 3467)	17.57*	14.64*	2.44	15.56*	2.57	14.29*
Family composition (FAM)						
Married couple	1186.09	1042.36	832.46	854.54	383.55	449.76
Married couple with 1 or 2 children	852.66	700.03	859.24	588.65	398.46	306.04
Married couple with 3 or more children	802.61	795.73	1211.86	637.12	510.47	337.55
Single person	892.12	710.29	771.48	603.69	322.62	300.02
Single parent with children	497.47	429.51	679.96	354.94	310.72	178.45
F-statistic (degrees of freedom: 4466)	3.00**	3.00**	2.40**	2.91**	1.95	3.09**

* Significant at 0.01 level
** Significant at 0:05 level
Significance indications based on two-tailed tests

fact that about half the families in the lowest income class had family heads of at least 65 years of age. The average length of stay for hospital separations for members of the oldest families was also over twice the sample average.

With respect to days' stay by age of family head, there is no significant difference when the entire sample is considered. When attention is restricted to families with at least one separation, a significant direct relationship appears between hospital use and age of family head.

Not surprisingly, the frequency of hospital use (separations) and total hospital use per family increases with family size. This is purely an aggregation effect, since on a per person basis it is the single-person family that is the greatest beneficiary of the insurance program.

As for the benefit incidence results based on our six different cost weightings, for the user subsample, average hospital benefits vary inversely with family income and directly with the age of the family head. Families consisting of at least a married couple are on average greater beneficiaries than single-person or single-parent families. Based on the most specific unit of observation - average benefits per person only for persons with one or more separations - the greatest beneficiaries are user members of 'married couple' families.

In general, the increasingly specific cost weight focus (from province-wide per diem to marginal inpatient hospital-specific and diagnosis-specific costs) does not alter significant differences or ordinal rankings. However, a number of interesting inferences and findings were obtained by the use of the alternative cost weightings. For example, families in the lowest income class were more likely to be treated in lower-cost hospitals than the rest of our sample. Families with heads in the 35-49 age bracket, on the other hand, were treated in higher-cost hospitals, as were members of the largest OHIP families. Single persons tended to be treated in lower-cost hospitals.

Another implication of the analysis is that the differences in use do not seem to be a function of differences in diagnosis. Only by applying inpatient and diagnosis-specific cost weights could we eliminate diagnostic mix or mix of hospital activities as potential contributing factors. The results based on diagnosis- and hospital-specific weightings indicate significant differences in benefits across income classes, with the lowest income group being the greatest beneficiaries, but the significance is less

marked than for the differences in incidence of use.

In Chapter 2 we reviewed several hospital use studies of the National Health Service in the United Kingdom, the Medicare and Medicaid programs in the United States, and the Canadian program. In this section we compare very briefly some of our findings with that earlier evidence.

As we noted earlier no US or UK study has incorporated monetary measures of hospital benefits such as those developed for this study. The studies by Boulet and Henderson (1979) and Statistics Canada (1977) include one measure of hospital benefits as explained in Chapter 2. That measure is different from any of the ones we have used here. It is closer to the CPDOW measure than to the other benefit measures in our analysis. For several reasons, a straightforward and complete comparison of the benefit magnitudes in the two studies is difficult. The most important of these is the unique definition of the family adopted in our study. Also, our focus was restricted to hospital services provided in acute-care hospitals, unlike the broader inclusion in the Statistics Canada data base. Perhaps most significantly, some of the more interesting results in our analysis in this chapter are directly derived from our different weightings of the utilization data. For example, our comparison of benefit incidence using province-wide and hospital-specific per diems as weights led us to the conclusion that the lowest-income families seem to be treated in relatively low-cost hospitals.

With these restrictions and caveats in mind, it can be said that our results pertaining to days' stay and average length of stay by income class, age of the head of the household, and family size basically concur with the Canadian studies. They also conform broadly to the results of studies in the United Kingdom, particularly the one by Carstairs and Patterson (1966), if one is willing to assume that there is a reasonable correspondence between our categorization of families on an income scale and the social class grouping normally used in the UK studies. The poor and the aged are found to be greater users of hospital care and to stay longer in hospital. Larger families are greater users of hospital care, though on a per capita basis it is the smaller families that enjoy this distinction.

The studies by Statistics Canada and Boulet and Henderson found that average hospital benefits vary inversely with family income (see Tables 8 and 9). Our results are again consistent with that finding, both

on a per family basis (Table 13) and on a per capita basis (Table 15), and regardless of whether the benefit measure used is based on CPDOW - which is close to the Statistics Canada measure - or on CPDHS.

It was noted earlier that the results presented in this chapter provide, at best, impressionistic evidence regarding the relative effects of selected socioeconomic variables on the variations in hospital use and benefits. In the next chapter we consider in more detail the relative influence of these and other explanatory variables. Because of the problems, noted in detail earlier, associated with using the case-based benefit measures (IPCHS and MCC) and in keeping with the focus of our discussion in this chapter, the benefit measures used in the causal analysis in Chapter 5 will be based on the CPDOW, CPDHS, IPDHS, and MCD cost weights.

5
SOCIOECONOMIC FACTORS IN THE INCIDENCE OF HOSPITAL BENEFITS

The results of our incidence analysis suggest that poor families are the primary users of hospital care. They also show that the use of hospital care and the benefits from hospital insurance increase with the age of the household head, whether the unit of comparison is the individual or the family. The largest families were, not surprisingly, found to be the most extensive users of hospitals, while on a per person basis, single-person families received the most hospital benefits. We surmised that the poor receive relatively more services than their counterpart families of other income classes because of the different age structures of the various income classes.

But while the main purpose of Chapter 4 was to estimate the average hospital use and benefits for each of our income classes, one cannot infer anything from those results about the role of income in explaining the variation in benefits between income classes. In this chapter, then, quite different questions are addressed. First, and most important for the issues being considered by this study, is family income itself a significant determinant of use of hospital services? Second, are there other variables of significance to government policy that appear to have a direct influence on patterns of hospital use and benefit? Third, and directly related to the first two questions, is the determinant role of other variables such as to suggest a significant difference between the apparent income/benefits pattern of Chapter 4, and the causative impact of income on the variation in benefits? Finally, what is the best analytical means of answering the first three questions?

There is a vast literature on general correlates of medical and hospital care, and somewhat more restricted but extensive research on the determinants of medical and hospital use. Here we draw only selectively on that past work in formulating our determinant model.

Perhaps the pioneering work on the taxonomy of the correlates of the use of medical services is by Andersen (1968, 8), who suggested that all factors fall into one of three groups - 'predisposing', 'enabling', and 'need'. He postulated that 'for use to take place: (1) a family must be predisposed to receive medical care; (2) there are enabling conditions that allow the family to attain health services; and (3) the family must perceive a need for these services'. Predisposing variables include family composition, social class (as determined by location, occupation, education, race, ethnicity, and so on), and health beliefs (related to families' sense of the value of good health, knowledge of disease, and the role of providers in maintaining good health). Enabling variables include the accessibility of services and the ability of families to afford those services. Income, insurance coverage, and direct costs fall into this category. Finally, the need variables are related to perceived health status and the response to symptoms. Aday and Eichhorn (1972) compiled an extensive bibliography (now somewhat dated) of studies examining the roles of various subsets of these variables; a more recent bibliography may be found in Freeburg et al. (1979).

Causal models of use may be classified in two ways. First, they can generally be divided into macro- and micro-analyses, depending on the unit of comparison. Macro-analyses use characteristics of populations to try to explain differences in use between regions or countries. A number of economic approaches to modelling hospital use are reviewed in Horne (1978). One more recent example is Bridgman (1979). Micro-analyses, on the other hand, focus on the family or individual as the unit of comparison and attempt to explain variation by using variables that describe the family or its members.

The second way of classifying causal models is to determine which groups of the many variables in Andersen's (1968) model are included in the analysis of use. McKinlay (1972) and Anderson (1973b) have reviewed the approaches taken by different disciplines (e.g. economic, social-psychological, and so on), while selective reviews of the economic approaches may be found in Berki (1972), Stoddart (1975) for medical rather than hospital use, and Horne (1978). Newhouse and Phelps (1976) and Horne (1978) are recent examples of micro-analytic economic approaches to explaining hospital use in mixed and public insurance settings respectively.

MODELLING HOSPITAL USE

Like Newhouse and Phelps (1976), Horne (1978), and Beck and Horne (1980), we treat use or non-use and the extent of use separately and employ a two-stage model for estimating the determinants of hospital use. The first stage investigates the influence of income and other independent variables on variations between families in the probability of at least one admission during the study period. The second stage then considers how the same set of independent variables explains the volume of hospitalization among 'user' families and the benefits to those families.

The division of the causal analysis into two parts is justified for two reasons, one statistical, the other intuitive. First, as discussed below, the large proportion of families without any hospital use during the period of analysis imposes a distribution of hospital use by family with a heavy weighting at zero. This distribution is statistically problematic if one wishes to explain the variations in hospital use between families for all families. One method of avoiding that problem is to deal with the two stages separately. Secondly, there is no reason to believe that the relative importance of any given independent variable will be the same in explaining the probability of admission and the length of stay once admitted. Yet there are different policy questions relating to each, and it seems logical, therefore, to incorporate the possibility that different variable-specific effects will emerge in each stage. For example, one might expect that a single-person family would be less likely to experience at least one hospitalization during a given period, other things being equal, than larger families. Similarly, because the family consists of a single person, one would expect relatively low use and benefits on a family basis. But the effect of that family composition could be different in the two stages, since that single person, once admitted, might be discharged with less haste than a similar patient who could be cared for at home by other family members. In other words, if all else, including type and severity of condition is equivalent, a single person, once admitted to a hospital, might stay longer than a person from a larger family. This would tend to generate coefficients of different magnitude (if not direction) on the same variable in the two stages of use.

We proceed, then, from the premise that similar variables have different weights in the two stages. The implications of this hypothesis

for choosing an estimation technique for the utilization model are discussed in more detail later in the chapter. In addition, the hypothesis itself is tested on our data set. It is assumed that the probability that a given family will need at least one episode of hospital care in any given period is a function of numerous economic, social, and demographic variables. Included are those specifically reported in the incidence results of Chapter 4, along with the education level and employment status of the household head. Variables are also included to identify families carrying supplemental hospital insurance, and family heads or spouses whose personal physician is a specialist. Finally, 'enabling' variables representing the availability of medical and hospital services complete the set of independent variables.

Our choice of independent variables was generally in keeping with other micro-economic analyses of hospital use,[1] but it was also intended to be as consistent as possible with the earlier analysis of the use of medical services for the same sample (Manga 1978), and was restricted somewhat by the availability of data. Table 18 lists the acronyms and descriptions of all the variables used in the model.

This causal analysis is restricted to measures of use and benefit that are based on days' stay as an output proxy because of the possibility that severe bias may be imparted to results based on cases by the fact that per case costs assume that lengths of stay (for given diagnoses and/or hospitals) are constant for all socioeconomic classes. Thus,

$$P(ADM) = \Phi(a_0 + a_1 INC + a_2 FAM + a_3 HAGE + a_4 BDPOP + a_5 DRPOP + a_6 HEDU + a_7 REMP + a_8 SDOC + a_9 SCOV), \qquad (1)$$

1 For example, Newhouse and Phelps' (1976) analysis of individuals included a number of price-related variables not relevant to our setting, as well as income, family size, education, age, sex and race, self-perceived health status and disability days, ratio of beds to population, ratio of physicians to population, and a rural/urban variable. Horne's (1978) analysis using the family as the unit of observation includes variables to distinguish family composition, welfare recipients, age of the household head, family size, and income. Our analysis does not explicitly include health status or need variables, because the survey did not collect such information. It is not clear to us, however, that this is any worse than using ex post self-reported perceived health status or disability days to proxy ex ante health status.

TABLE 18

Variables used in use and benefit equations

Dependent variables

P(ADM)	1 if any member of family was admitted to hospital during year; 0 otherwise
DAYS	Total days of hospitalization for all family members
BFT_i	Total hospital benefits for all family members based on cost estimates:

i = 1 - CPDOW
2 - CPDHS
3 - IPDHS (as described in Chapter 3)
4 - MCD

Explanatory variables

INC - family income

INC1	1 if \$0 ≤ INC ≤ 3999; 0 otherwise
INC2	1 if \$4000 ≤ INC ≤ 7999; 0 otherwise
INC3	1 if \$8000 ≤ INC ≤ 13,999; 0 otherwise
INC4	1 if \$14,000 ≤ INC ≤ 19,999; 0 otherwise
INC5	1 if INC ≥ \$20,000; 0 otherwise
INC6	1 if family income is not known; 0 otherwise

FAM - OHIP family composition

FAM1	1 if a married couple with no children; 0 otherwise
FAM2	1 if a married couple with one or two children; 0 otherwise
FAM3	1 if a married couple with three or more children; 0 otherwise
FAM4	1 if one person only; 0 otherwise
FAM5	1 if a single parent with children; 0 otherwise

HAGE - age of head of household

HAGE1	1 if 18 to 34 years of age; 0 otherwise
HAGE2	1 if 35 to 49 years of age; 0 otherwise
HAGE3	1 if 50 to 64 years of age; 0 otherwise
HAGE4	1 if 65 years of age or more; 0 otherwise

BDPOP - acute care beds per 10,000 population in county where family lives

DRPOP - practising physicians per 1000 population, in county where family lives

HEDU - education of head of OHIP family

HEDU1	1 if not passed high school; 0 otherwise
HEDU2	1 if only graduated from high school; 0 otherwise
HEDU3	1 if has technical training or some college or university education; 0 otherwise
HEDU4	1 if graduated with university degree; 0 otherwise

TABLE 18 continued

REMP - employment status of head of family	
REMP1	1 if self-employed or employed full-time; 0 otherwise
REMP2	1 if employed part-time, laid off, or unemployed; 0 otherwise
REMP3	1 if retired or disabled; 0 otherwise
REMP4	1 if full-time student or housewise; 0 otherwise
SDOC	1 if family members have a regular personal physician who is a specialist; 0 otherwise
SCOV	1 if family has supplemental hospital insurance coverage; 0 otherwise.[a]

a A family is assigned an SCOV value of 1 if its respondent answered 'yes' to either of questions 5(b) or 5(c) (see Manga 1978, 168).

where Φ (x) is the cumulative normal function (see appendix);

$$DAYS = b_0 + b_1 INC + b_2 FAM + b_3 HAGE + b_4 BDPOP + b_5 DRPOP + b_6 HEDU + b_7 REMP + b_8 SDOC + b_9 SCOV; \text{ and} \tag{2}$$

$$BFT_i = d_{0i} + d_{1i} INC + d_{2i} FAM + d_{3i} HAGE + d_{4i} BDPOP + d_{5i} DRPOP + d_{6i} HEDU + d_{7i} REMP + d_{8i} SDOC + d_{9i} SCOV. \quad (i = 1, \ldots, 4) \tag{3}$$

Each independent variable, with the exception of BDPOP and DRPOP, actually represents a set of dummy variables.[2] Thus, INC denotes income class and comprises the six-class breakdown used in Chapter 4. The set of families contained in each cell for each variable is outlined in Table 18.[3]

The dependent variable for equation (1) represents the first stage of the utilization process - probability of at least one family admission to hospital. Equation (2) is the utilization segment of the model, while equation (3), which is really four separate equations, comprises alternative specifications of the hospital benefits model. The values of all variables

2 BDPOP and DRPOP were the only independent variables for which data sources allowed us to create continuous variables. All other such variables were derived from survey responses.

3 Of course one cell from each set of dummy variables is dropped at the model estimation stage to avoid collinearity. See Manga (1978, 88-90) for a justification of the use of dummy variables.

are for the twelve-month study period.

The following section discusses the conditions under which equation (1) would be redundant (i.e. if we could not reject the hypothesis that the same variables have the same weights in each stage of the utilization process). Before turning to that issue and to other analytic questions, we consider briefly some plausible expectations about signs and ordinal rankings within groups of variables for the model's parameters.

When there is universal public hospital insurance and essentially no direct out-of-pocket costs for hospital care,[4] the traditional 'enabling' role of income disappears. Unless there are less direct influences, one might reasonably expect income class to have no significant impact on the distribution of hospital benefits. That would not be inconsistent with the results presented in the previous chapter, for benefit incidence is indicative only of correlative patterns.

Income has been postulated to affect the use of services in a number of 'non-enabling' ways. It has been argued that wage income is inversely related to use because of the opportunity cost of lost work time (Grossman 1972). If we come at this from the other end, hospitalization itself may reduce income (Luft 1975), particularly among those without sickness and disability benefits. Income may also interact with education since 'high income families tend to be better educated, and hence might well be more cognizant of the benefits to be derived from early detection of illness and/or maintenance of good health' (Horne 1978, 135). In short, there are indirect reasons for believing that family income might be inversely related to hospital use and benefits in the setting for our analysis.

To the extent that the age of the family head is a proxy for the age structure of the entire OHIP family, this variable is primarily a 'predisposing' variable. But because the ability to perceive illness and a family's reaction to symptoms are likely to change with aging, the variable also has a 'need' component (Andersen 1968). In any event, one would expect that a family's age structure would be an important determinant of its use of hospitals. In particular, we might expect the HAGE4 coefficient to be the largest of the group, followed by either HAGE3 or HAGE1, depending on the extent of the maternity influence. In particular, HAGE1

4 Indirect costs may, of course, remain. Travel time and costs and queuing time will not be equally distributed by region, but their distribution by income class is unknown.

would be expected to be relatively more important in the probability of admissions equations.

Similarly, family composition will reflect not only a family's stage in its 'life cycle' (Horne 1978, 124), but also the availability of support in the family. While by reason of sheer numbers one expects the probability of hospitalization to increase with family size, it may also be reasonably argued that, for any given illness requiring hospitalization, members of large families will incur shorter hospital stays than members of smaller families. Thus, while hospital use would be expected to increase with family size, we would not expect a linear relationship. Families in the FAM3 group would be likely to rank highest in both probability of admission and in use and benefit, and FAM2 members would probably be second. The ordinal relationship of FAM1 and FAM5 is difficult to guess, but FAM4 might be expected to show the lowest parameter in all equations.

The employment status of the household head is included as a proxy for a number of potential contributing influences. As noted above, part-time, unemployed, or student status reflects a low opportunity cost in lost wages, although sickness and disability benefits may mean a low wage opportunity cost for those fully employed as well. Depending on the type of work, 'professional/career advancement' opportunity cost may be a more significant factor than short-term lost wages. 'Retired or disabled' proxies a component of health status as well as reflecting low opportunity cost. On this premise, one might expect lower hospital use by persons working full-time than by the unemployed, and particularly high use by the retired or disabled.

The effects of education level on hospital use are cloudy at best. For medical services it has been argued that 'the more educated demand fewer episodes [of treatment] because they are also likely to possess more information about health matters that allows them to [treat themselves, but] ... it is at least equally likely that the more educated have a greater awareness of which symptoms require attention in addition to their likely greater knowledge of the availability of medical services' (Stoddart 1975, 86). In fact, even the direction of causality is unclear (see Manga 1978 and Grossman 1975). As for the links between education and hospital care, for conditions severe or complex enough to require hospitalization, relative ability to treat oneself is probably not a relevant consideration. On the other hand, the better-educated may be more aware, not only of

symptoms signifying a need for hospitalization, but also of the inherent dangers of hospital treatment itself. The jury is still out.

The supply of beds tends to be highly correlated with bed use. That is not sufficient justification, however, for including a bed supply variable among the independent variables in a model of this type. As Horne (1978, 23) notes, its inclusion rests not on the hypothesis that 'bed supply represents (as it ultimately must) a constraint on bed use, but rather that bed supply is itself a key determinant of bed use.' Since our setting precludes a 'supply increase → price fall → increase in quantity demanded' chain of events by reason of the absence of direct charges (not to mention the potential role of the physician), we can test for a 'pure availability effect' (Feldstein 1971), which is 'produced in non-standard fashion by a rightward shift in the supply curve which induces a corresponding rightward shift in the demand curve' (Horne 1978, 24). When there are no direct price effects, 'the proposition that the supply of beds creates its own demand' is not ambiguous and can be tested without 'a demand function for hospital care in which both price and availability are explanatory variables' (Feldstein 1971, 30-1). In point of fact, however, the very suggestion of a pure availability effect implies that one would be estimating a utilization function, not a demand function (Stoddart and Barer 1981). After all, at least some non-price-reaction effect will work through the physician's role in determining admission and length of stay. At any rate, a positive coefficient on the BDPOP variable would be consistent with an operative pure availability effect.

There are numerous stories one could tell about the effect of the supply of physicians on hospital use. It could be argued, for example, that where there is a low density of physicians patients do not make contact with doctors as often and therefore will be hospitalized less often for conditions discovered in the course of routine investigations. Alternately, one could suggest that a high density of physicians makes non-inpatient treatment of routine conditions or investigative work-up more prevalent than in less urban areas. It has even been suggested that physicians at times admit or retain patients in hospitals in order to 'protect' their allocation of admitting privileges (Harris 1977).

It is difficult to predict which effect(s) will dominate, but a positive relationship between BDPOP and use would lend support to the first and last suggestions.

Finally, since a large proportion of hospital admissions are channelled through specialists, families with a regular specialist might be expected to incur greater hospital costs than their counterparts that have no family doctor or consult a general practitioner. We would not expect SCOV to have any significant impact on the probability of admission, unless having supplemental hospital insurance is viewed as an ex ante health status proxy. If the families with higher expectations of a need for hospitalization tend also to be the most frequent purchasers of supplemental coverage, and if those families subsequently require more hospitalization than their socioeconomic peers, then we would of course expect a positive parameter estimate in all equations. In addition, once admitted, a patient in a ward bed may exert greater pressure to be discharged than a similar patient in a semi-private or private room. A more significant positive coefficient might therefore be expected in equations (2) and (3) than in equation (1).

ISSUES IN MODEL ESTIMATION

Although most families receive some medical care each year, hospitalization is a comparatively rare event. As a result, any random sample of families will include many with no hospitalization. Where hospital use and benefits form the dependent variables for a model such as the one used here, those dependent variables will often take on the value of zero.

Of the 1290 families constituting our data set, less than 30 per cent incurred any hospital costs. From a model estimation perspective, this situation clearly violates the assumptions of normal least squares regression, (that is, normally distributed errors and homoscedasticity).[5]

Accordingly, it was necessary to use a refinement of the usual model, designed to handle data sets characterized by a frequency distribution with a concentration point at zero. Here we discuss briefly the various methods considered. Analytic detail is confined to the appendix.

5 Note that it is the almost 70 per cent of families with no hospitalization, rather than the 'either-or' nature of hospitalization, that undermines the use of ordinary least squares estimation. If every family experienced at least one hospital separation, hospitalization would be no less an 'either-or' event (in the sense that negative values for the dependent variable would be precluded), but the distribution of days' stay could be fairly normal.

One possible method of estimation is Tobit analysis (Tobin 1958). An example of its use may be found in Davis and Reynolds' (1976) analysis of access to medical and hospital care under Medicare and Medicaid in the United States. Tobit analysis assumes an underlying normal distribution and then replaces all negative values with zeros. Parameter estimation and significance testing are conducted using maximum likelihood estimation. For the present study, this model was considered and found wanting (see appendix). Instead, a two-stage estimation procedure was adopted, for reasons outlined in the previous section. In particular, the Tobit model would impose equivalent parameters for the admission and stay stages of hospitalization, whereas the two-stage method allows an added dimension of flexibility.

The first estimation stage relates our independent variables to the probability of admission (equation 1). The second stage considers the impact of those same independent variables on the extent of use and the amount of benefits for families that incur some hospitalization. Implicit in this formulation is the assumption that the covariates (set of independent variables) that are important in determining admission may not have the same effect on hospital use for families with admissions. The difference may be in direction or magnitude.

In the appendix, two different two-stage estimation techniques are discussed - Probit plus truncated normal; and Probit plus ordinary least squares. For theoretical and analytic reasons, outlined there, the latter method was adopted and the results of the next section are based on it.

QUANTITATIVE RESULTS

Stage one - Probit analysis

The Probit model assumes that covariates affect the probability of admission as described in equation (1). The parameter estimates are presented in Table 19. To test for the significance of INC, for example, one compares the full model to a model that omits the set of INC dummy variables. In this analysis it was found that after we had adjusted for other covariates, the family income variable is not important in predicting admission ($\chi^2_{(5)} = 3.14$). The significant covariates were FAM1, FAM4, DRPOP, and REMP1. This implies that single-person families (FAM4) and

TABLE 19

Parameter estimates from Probit model: regressing the probability of admission on selected covariates using the whole sample (n = 1290)

Covariate	Coefficient	t-statistic
Intercept	-0.12	0.30
INC1	-0.20	0.99
INC2	-0.29	-1.54
INC3	-0.17	-1.06
INC4	-0.23	-1.35
INC5	-0.25	-1.47
FAM1	-0.45	-2.37**
FAM2	0.07	0.39
FAM3	0.23	1.21
FAM4	-0.78	-3.90*
HAGE1	0.36	2.00
HAGE2	0.19	1.06
HAGE3	0.23	1.44
DRPOP	-0.18	-2.57**
BDPOP	0.06	1.70
HEDU1	-0.04	1.33
HEDU2	0.05	0.36
HEDU3	-0.12	0.92
SDOC	0.18	1.29
SCOV	0.10	1.25
REMP1	-0.44	-2.32**
REMP2	-0.28	-1.12
REMP3	0.30	1.43

* significant at 0.01 level
** significant at 0.05 level
$\Delta\chi^2_{(5)}$ for income = 3.14

couples without children (FAM1) would have a significantly lower probability of admission than families of other compositions when all other determinants of admission are taken into account. This concurs with what was suggested by the cross-tab results in Chapter 4 and our discussion earlier in this chapter. Similarly, it is reasonable for families headed by a self-employed or fully-employed adult (REMP1 families) to have a lower probability of admission. Such families are likely to have better health status and greater opportunity cost of illness than families in the other employment categories. We argued earlier that the relationship between the supply of physicians and the probability of at least one admission could be positive or negative. We found the effect to be negative and surprisingly significant. While we are unable to offer any definitive

explanation, this effect may reflect a phenomenon noted earlier - where there is a high density of physicians, patients come in contact with physicians more frequently than elsewhere (Manga 1978), and consequently conditions that would eventually require hospitalization if not attended to may be detected sooner and treated on a non-inpatient basis.

While no other covariate proved to be significant, we note that both the sign and relative magnitude of the coefficients in Table 19 generally conform to our expectations. Thus the signs on both the FAM2 and FAM3 variables are positive and the coefficient on FAM3 is the larger of the two. Similarly, both SDOC (family members have a regular personal physician who is a specialist) and SCOV (family has supplemental hospital insurance coverage) are insignificant determinants of probability of admission.

The fact that income is not significant implies that when account is taken of all other determinants of probability of admission, income itself does not influence this probability. If there is an opportunity cost effect not captured by REMP, there is less reason to expect it to influence the probability of admission than length of stay once admitted, because the opportunity cost will increase with time away from productive employment. There is a positive relationship between bed supply and admission, but it was found to be insignificant. For reasons explained above, we had no expectation about the effect of education on admission. The Probit regression reflects this unclear and insignificant relationship between the two variables.

The decision to admit a patient to hospital is almost exclusively made by a physician, and this decision is presumably made on the basis of medical need. The nature and course of treatment, too, are largely dictated by the specifics of the medical problems of the patient. However, to some extent, location of treatment (hospital or non-hospital) and particularly of convalescence may be influenced more by the economic and social situation of the patient than was the admission decision. Many factors may be relevant. Does the patient get sick-leave? Is the hospital a more pleasant place than the patient's home? Does the patient have relatives to care for him or her at home? Does the patient have access to home care, a nursing home, or other long-term care institutions? These and related questions often determine whether a patient spends more time in hospital than seems necessary for purely medical reasons.

Stage two - ordinary least squares

To analyse the variation between families in hospital use and benefits, we have specified two statistical questions. First, how is the probability of admission related to the socioeconomic and demographic characteristics of those families? Second, how is the variation in use and benefits related to the same socioeconomic and demographic variables?

The data used to address the latter question contain no negative values. The benefit and use variables are more likely to be sufficiently continuous and the error term sufficiently normal that one may safely use the ordinary least squares technique of estimation on the subsample of families with one or more admissions. Normal linear regressions were performed on each of equations (2) and (3) (five equations in all) for the 377 families with admissions. The parameter and t-statistics appear in Table 20.

One of the most striking pieces of information in this table is the magnitude of the R^2's. Cross-sectional analyses of this type, which look at use and expenditures in the health sector and include no personal health status measures, have traditionally reported low values for R^2 or for other goodness of fit measures (Beck 1974, Manga 1978, Acton 1975, Davis and Reynolds 1975, 1976, Horne 1978). On the other hand, although the R^2 values of Table 20 show that a good deal of the variance remains to be explained, they are relatively high values for this kind of cross-sectional study.

The equations using dependent variables based on CPDHS, IPDHS, and MCD generate marginally higher R^2 values than those based on days' stay (or CPDOW). This seems to suggest that the refinement of using hospital-specific cost measures improved upon the use of province-wide per diems but that moving through inpatient- and diagnosis-specific modifications did not generate additional explained variance.

We turn now to the impact of groups of variables. Table 21 indicates the groups that were significant; that is, the groups of variables for which variance within groups was a significant determinant of the variance in each of our five dependent variables. In our analysis of different estimation techniques, we used OLS and Tobit methods to estimate parameters for the whole sample (n = 1290). In both cases income turned out to be an insignificant explanatory variable. As Table 21 shows, the

TABLE 20

Coefficients from ordinary least squares regressions for families with at least one admission (n = 377)

	Measures of hospital use or benefits								
	Days' stay		BFT (CPDOW)[a]	BFT (CPDHS)		BFT (IPDHS)		BFT (MCD)	
Variable	Coeff.	t-statistic	Coeff.	Coeff.	t-statistic	Coeff.	t-statistic	Coeff.	t-statistic
Constant	-14.9	-1.7	-1574	-1856	-2.2**	-1470	-2.2**	-762	-2.0**
INC1	15.1	3.4*	1601	1251	3.0*	1059	3.1*	438	2.3**
INC2	7.4	1.8	781	758	1.9	618	2.0**	319	1.9
INC3	8.7	2.5**	920	956	2.9*	753	2.8*	405	2.7*
INC4	2.6	0.7	279	453	1.4	336	1.2	157	1.1
INC5	4.0	1.1	418	434	1.3	346	1.2	174	1.1
FAM1	-2.9	-0.9	-312	-280	-0.9	-237	-0.9	-109	-0.9
FAM2	2.5	0.9	268	122	0.5	124	0.6	79	0.7
FAM3	6.8	2.1**	718	762	2.6**	602	2.5**	297	2.2**
FAM4	-9.5	-2.8*	-1005	-1126	-3.6*	-880	-3.5*	-468	-3.3*
HAGE1	2.9	0.7	307	547	1.4	408	1.3	210	1.7
HAGE2	0.3	0.1	35	422	1.1	288	1.0	222	1.4
HAGE3	5.0	1.3	528	791	2.2**	607	2.1**	354	2.3**
DRPOP	3.1	2.0**	327	479	3.4*	365	3.2*	141	2.2**
BDPOP	0.4	0.6	42	51	0.8	39	0.7	22	0.7
HEDU1	2.7	1.0	288	192	0.7	168	0.7	79	0.7
HEDU2	2.8	0.9	300	264	0.9	228	1.0	116	0.9
HEDU3	2.5	0.9	269	190	0.7	167	0.7	113	0.9
SDOC	5.1	1.9	538	521	2.0**	405	1.9	303	2.6*
SCOV	4.1	2.4**	437	322	2.0**	281	2.2**	134	1.9
REMP1	2.9	0.6	311	22	0.1	74	0.2	38	0.2
REMP2	-3.2	-0.5	-334	-373	-0.9	-282	-0.6	-176	-0.7
REMP3	19.5	3.8*	2067	1947	4.0*	1594	4.0*	929	4.3*
R^2	0.269		0.269	0.283		0.284		0.282	

* Significant at 0.01 level

** Significant at 0.05 level

a The t values for the benefit equation using CPDOW weights are, of course, identical to those for the days' stay equation

TABLE 21

F values for each variable group for OLS regressions of Table 20 (n = 377)

	STAY CPDOW	CPDHS	IPDHS	MCD
INC (5,354)	3.93*	3.10*	3.29*	2.84**
FAM (4,354)	9.58*	13.07*	12.64*	11.27*
HAGE (3,354)	1.87	2.26	2.25	2.04
HEDU (3,354)	0.34	0.27	0.31	0.34
SDOC (1,354)	3.45	4.04**	3.76	6.95*
SCOV (1,354)	5.92**	4.01**	4.71**	3.55
DRPOP (1,354)	4.20**	11.24*	10.06*	4.95**
BDPOP (1,354)	0.34	0.59	0.54	0.58
REMP (3,354)	12.04*	16.00*	15.70*	17.78*

NOTE: degrees of freedom are in parentheses
* Significant at 0.01 level
** Significant at 0.05 level

application of such estimation techniques to samples not well suited to the assumptions inherent in their use can (and in this case does) generate misleading results. In fact, Table 21 is interesting in two respects for the income variable. First, when hospital use is viewed as a two-stage process and attention is focused on the experiences of families that have members admitted to hospital, we see that income is a significant determinant of variance in hospital use and benefits. Second, as the F-statistics for income illustrate, the degree of confidence we may have in rejecting the null hypothesis that income is not significant shrinks slightly when we move away from using the simple benefit measure based on CPDOW. At a

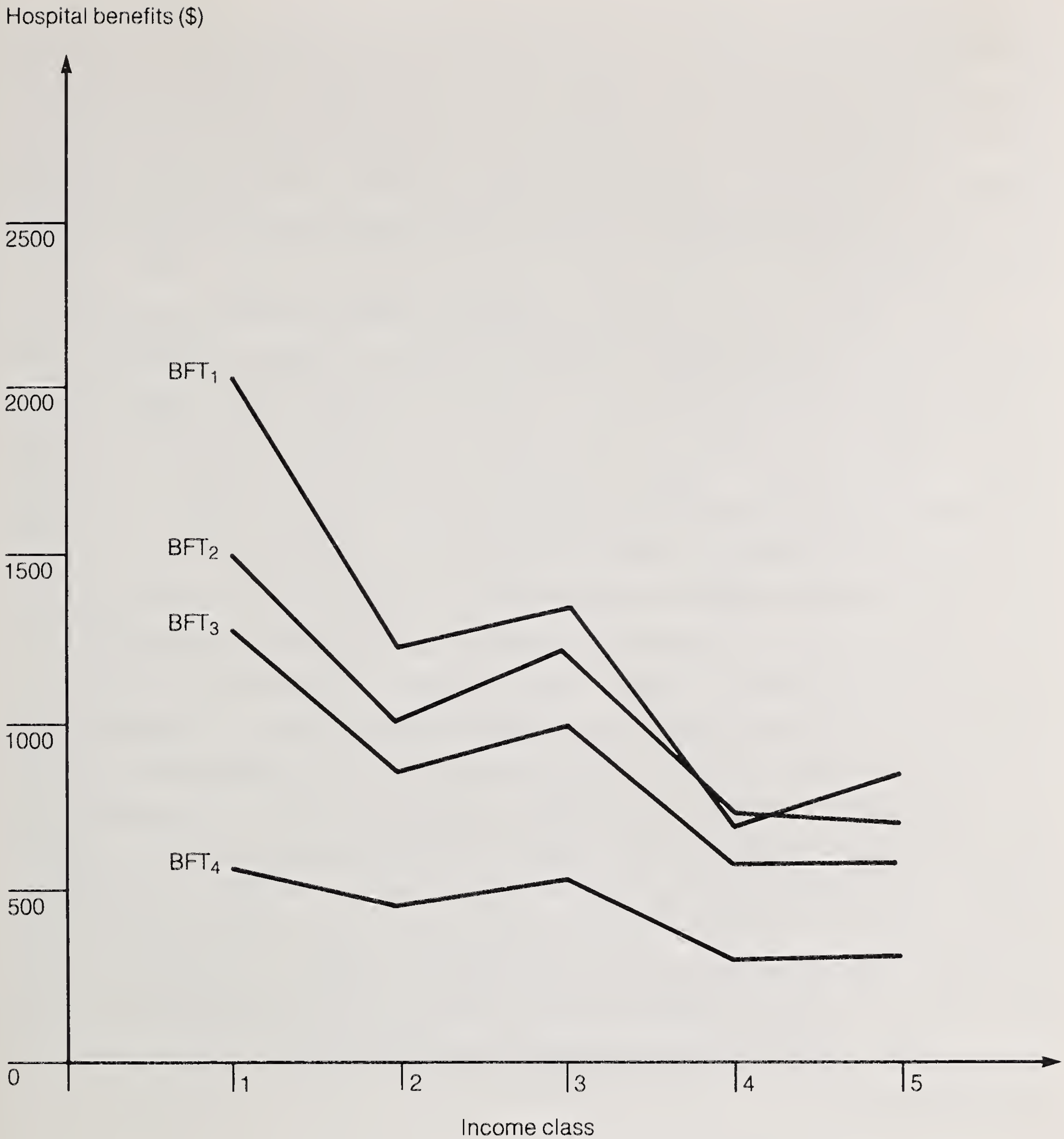

Figure 1
Estimated means of hospital benefits by income group using CPDOW, CPDHS, IPDHS, and MCD

99 per cent confidence level, one is able to reject the null hypothesis using the benefit measures based on CPDHS and IPDHS, while it can only be rejected with 95 per cent confidence for the equation based on MCD. Still, such a critical level suggests that income is important for the hospital users.

The shrinking significance of the income variable as one moves to more precise cost weights is illustrated also in Figure 1, where the estimated mean benefits are plotted by income class for each of the four benefit measures based on cost per day. These mean values were determined by setting the covariates other than income at their mean values and using the parameter estimates of Table 20 to generate income-class-specific mean benefits.

More detail on the effect of income is shown in Table 20. There it can be seen that the parameter estimates on the $0-3,999 and $8,000-13,999 income classes are significant ($p < 0.05$), irrespective of which use or benefit measure is used. Thus, being in either of these two income classes seems to be a significant factor in explaining variance in hospital use among families with at least one admission. Although this is not what we expected, there are a number of possible explanations. If one looks at the results for the HAGE variables, it can be seen that the coefficients there, all of which are positive (and the one for HAGE3 significant), do not concur with what we might have expected - negative and significant coefficients at least on HAGE1 and HAGE2 since HAGE4, the aged group, is dropped. Instead, those results suggest that the hospitalized families with the oldest heads will, other things being equal, use less hospital care than the families in HAGE1 and HAGE2, and significantly less than those in HAGE3. Also surprising is the result in Table 21 that, as a variable group, variance in age of head of household is not a significant factor underlying variance in hospital use among user families. Since it is known that a large proportion (around 40 per cent) of families in HAGE4 are also in INC1, the significant parameter on the latter may be indicative of that interrelationship.[6]

6 This correlation may raise questions about the absence of interaction terms in our analysis. However, correlation and causal interaction are two different phenomena. In the extreme, for example, one might have almost every family with a head at least 65 years old also being in the lowest income class (i.e. nearly perfect collinearity). This certainly does not logically require a <u>different income effect</u> for older families in that income class than for all other families in the same income class, although that may be the case. Nevertheless, the reason for the lack of interactive analysis in this study was insufficient sample size. We could think of no a priori basis on which to choose a few interaction terms from among the many possibilities. Although the inclusion of large sets of these terms would not, strictly speaking, have absorbed all our degrees of freedom, we could

In addition, the strength of the INC1 parameter may simply reflect the lack of health status and severity of disease standardization. Although we have isolated the families that use hospitals, and built in case-mix cost variance, it is possible that, for any given diagnosis, the poor need hospital treatment for cases of greater severity. This may be simply because they let symptoms progress further before seeking treatment, or because when they do get sick, they have weaker recuperative powers than members of other income classes.

As for the significance of the parameter estimate on INC3, the only explanation we can offer is that this income group includes a disproportionate number of families with women in their peak fertility years who, with their children, use hospitals relatively often. Manga (1978) found that on a family basis this income class had the highest expenditures related to birth. Our Probit results in Table 19 show INC3 to have a larger coefficient (although not significantly so) for probability of at least one admission than the other four income groups.

We turn now to family composition. Because of the aggregative effect of family size on per family hospital use, we expected the parameter for FAM3 to be positive and significant, FAM2 to show the second largest parameter, and FAM4 the lowest parameter. This, not surprisingly, is borne out by the results in Table 20. As Table 21 shows, family size is a significant determinant of variation in hospital use and benefits for families that received some hospital care; and variation in use and benefits of the largest (FAM3) and smallest (FAM4) families in relation to the rest is the driving force behind the significance of that variable group.

The HAGE parameters are puzzling. Rather than negative and

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easily have created situations in which there were as many degrees of freedom in the error as in the model, which would have placed the stability of all our coefficients in doubt. In addition, the total number of observations for particular cells of the interaction variables becomes very small in many instances (e.g. for user families whose head is 65 years or older with income above $20,000). If one tries to avoid these problems by creating more aggregative interaction terms, then one faces yet another type of statistical problem - possible correlation among the interaction terms and the basic variables used in constructing them. Finally, while it is impossible to make direct inferences about hospital use from evidence about physicians' services, Manga (1978) found that interaction terms based on combinations of HAGE, INC, and FAM were always insignificant.

significant parameter estimates on HAGE1 and HAGE2 and at least a negative estimate on HAGE3, we see positive parameters on all three variables and 0.05 level significance on the parameter for HAGE3. One explanation based on the interaction of the HAGE and INC variables was mentioned above. Another and complementary explanation may be found in the strongly significant parameter estimate for REMP3 - the families with a retired or disabled family head. Again, since many REMP3 families are also HAGE4 families, the REMP3 parameter seems to be picking up some of the expected effect of HAGE4. The size and strong significance of the REMP3 parameter (which is reflected in Table 21 in the significance of the REMP group) may then be picking up some HAGE4 'noise' in addition to the expectations about opportunity cost. Interestingly, a similar pattern was found by Manga (1978) in his analysis of the use of physicians' services.

The lack of significance of the parameter on BDPOP suggests that we cannot reject the null hypothesis of no pure availability effect. DRPOP is somewhat more difficult to sort out. It was suggested that a positive coefficient would lend support to Harris's (1977) theory that physicians in areas with an abundance of physicians ensure that the bed complement 'assigned' to them is kept filled in order that it will not be reduced. Another possible explanation, that the positive and significant coefficient may reflect a causal chain running from greater availability of doctors to more frequent contact with doctors to more frequent discovery of conditions requiring hospitalization, seems ruled out by the negative coefficient on DRPOP in the probability of admission stage.

The earlier discussion of our expectations was unable to suggest a definitive direction for the HEDU variable (education of head of family). A number of possible links were noted along with the possibility that each might tend to cancel the effect of the other. Indeed, Table 21 shows that HEDU was not a significant determinant of variance in hospital use and benefits among families with admissions.

This leaves SDOC and SCOV to consider. We suggested that SDOC and SCOV would both show positive parameters for this stage of the analysis, as indeed they do. These are the only two parameter estimates that are significant ($p < 0.05$) in some equations but not in others. Thus SDOC is significant in the equations using benefit measures based on CPDHS and MCD, but not in the other three equations. In fact, the

coefficient becomes strongly significant (p <0.01) when the benefit measure based on MCD is the dependent variable, an indication that families that receive care regularly from a specialist tend to be hospitalized for more expensive conditions than average. Failure to use diagnosis-specific costs for benefit estimation will lead to underestimates of benefits for families with SDOC = 1. SCOV is significant until one gets to the benefits equation based on MCD; this perhaps reflects its role as a weak proxy for health status. In all cases the t-statistics hover around the critical (0.05) level, so that the flipping in and out of the 5 per cent significance range is not of paramount importance. It should also be noted that the relationship between SDOC and the hospital use and benefit measures may be 'associative' rather than one of cause and effect. That is, persons who are hospitalized may be more likely to consult a specialist regularly than those who are not, by reason of their need for hospitalization.

Before leaving this discussion of the interpretation of the OLS results, we return briefly to the income variable. Table 22 compares the use and benefit incidence results derived from the two methods used in Chapter 4 and the present chapter. It is an amalgamation of Tables 11 and 14 and Figure 1 (whose basis is Table 20). As noted earlier, the use and benefit incidence estimates from OLS were calculated by using the parameter estimates of Table 20 to compute expected days' stay and benefits for an 'average family' with a number of different incomes. Thus, a hypothetical family characterized by the mean values for all variables other than income (based on the 377-family user sample) was placed in each income class in turn, and the expected use and benefits were computed for that family.

Table 22 shows that the benefit incidence results are quite similar for the two methods. Even after we had standardized for the effects of the various socioeconomic covariates, the results by income class still show that the poor are the major users and beneficiaries. For the remaining groups there are only relatively slight changes in magnitude and order. But these results do show that some of the significant differences between income classes described in Chapter 4 were the result of income acting as a proxy for other socioeconomic variables. Irrespective of the dependent variable used in our regression analyses, the income group F-statistic is lower for the estimates based on regression analysis than for those deriving from the cross-tabulation analysis. Indeed, for the benefit equation based on MCD, the F-statistic drops below the 99 per cent significance level.

TABLE 22
Alternative hospital use and benefit estimates by income class for families with one or more discharges, using cross-tabulation and OLS regression analysis

Use or benefit measure	Type of analysis	INC1 ($)	INC2 ($)	INC3 ($)	INC4 ($)	INC5 ($)	F
BFT_1	CT[a]	2212	1396	1241	830	764	5.8*
(based on CPDOW)	OLS[b]	2065	1246	1385	743	883	3.9*
	Difference (CT-OLS)	+147	+150	-144	+87	-119	
BFT_2	CT[a]	1784	1255	1088	778	642	4.3*
(based on CPDHS)	OLS[b]	1548	1054	1253	750	730	3.1*
	Difference (CT-OLS)	+236	+201	-165	+28	-88	
BFT_3	CT[a]	1495	1036	895	635	536	4.7*
(based on IPDHS)	OLS[b]	1326	885	1020	603	613	3.3*
	Difference (CT-OLS)	+169	+151	-125	+32	-77	
BFT_4	CT[a]	744	541	485	322	278	4.0*
(based on MCD)	OLS[b]	590	471	557	309	325	2.8**
	Difference (CT-OLS)	+154	+70	-72	+13	-47	
DAYS' STAY	CT[c]	20.9	13.2	11.7	7.8	7.2	5.8*
	OLS[b]	19.5	11.8	13.1	7.0	8.3	3.9*
	Difference (CT-OLS)	+1.4	+1.4	-1.4	+0.8	-1.1	

* Significant at 0.01 level
** Significant at 0.05 level
a CT means cross-tabulation
b Estimated from OLS regression in Table 20
c Derived from Table 11

Comparison with previous Canadian evidence

Whereas our Chapter 4 results were compared with those of Statistics Canada (1977) and Boulet and Henderson (1979), those studies did not undertake a causal analysis of use. The past evidence that may be compared to our results in this chapter is restricted to two studies.

In a recent look at the patterns of hospital use by Saskatchewan families over the six years 1966-73, Beck and Horne (1980) do a similar

two-stage analysis that focuses on the effects of co-payment on the probability of admission and total use for users. Unfortunately, they report only the results of the probability of admission equations for their entire model. Since they include a continuous income variable, their model is restricted to identifying linear effects. They report a significant negative parameter estimate, which indicates higher probability of hospitalization for lower-income families. Horne (1978) earlier reported similar results. However, we found no significant effect of income on the probability of admission (Table 19).

Horne's analysis also investigated the effect of socioeconomic, demographic, and price variables on a number of dependent variables, including aggregate days' stay and diagnosis-specific days' stay. We are only able to compare our results with the former, more aggregative analysis. Again income was entered linearly, and Horne found a significant negative impact of income on days' stay per user family. As Figure 1 illustrates, our results agree in a very general way with those of Horne.

SUMMARY

The model of hospital use presented in this chapter is one that considers hospitalization as a two-stage process (probability of admission and length of stay once admitted). We have used Probit and ordinary least squares regression techniques respectively in the two stages.

It was found that the probability of having at least one family member admitted at least once to hospital was not affected by income. In fact, the only significant variables were FAM1 (married couple with no children, negative impact), FAM4 (single-person families, negative impact), DRPOP (physicians per 1000 population, negative impact), and REMP1 (self-employed or employed full-time, negative impact). Among the families that had some hospitalization, income, family composition, physician supply, and employment status were important explanatory variables irrespective of the formulation of the dependent variable, and SDOC and SCOV were significant in a majority of the five equations. Within variable groups, the marginal impacts were most often in the direction that had been postulated.

Table 22 illustrates two important points. First the straightforward cross-tabulation benefit incidence analysis of Chapter 4 typically overstated the benefits to the lowest income classes (INC1 and INC2) and understated

the benefits for the highest income group and INC3. In doing so it imparted a pro-poor bias to the differences between income classes in use and benefits for the user subsample, as reflected in the lower F-values for the OLS results in Table 22.

The second point relates to the application of different weightings for converting use to monetary measures. The effect of increasing degrees of specificity is shown vividly in Figure 1. It is evident that the differences in benefits between income classes are less pronounced when the benefits are based on diagnostic information than when either the average provincial per diem or the hospital-specific per diems are applied to days' stay. But BFT4 is not only the sole benefit measure incorporating diagnosis-specific information but also the only measure based on marginal rather than average costs. Thus the other three benefit measures imputed at least total inpatient operating costs to users of hospitals. When only the marginal costs are allocated to users, the differences in benefits between income classes become less significant. The use of average cost weights overallocates costs (benefits) to lower-income families with longer average lengths of stay. The diagnosis-specific weights for marginal costs also had the advantage of enabling us to determine that families that see a specialist regularly tend to be hospitalized for relatively expensive conditions.

Whereas a straightforward OLS analysis or Tobit analysis of our entire sample (n = 1290) would have suggested that income is not a causal factor in explaining differences in hospital use and benefits, our two-stage investigation suggests that it is unimportant to the probability of admission but that it is a significant determinant of use for the families with admissions. Since the sign of the parameter estimates shows clearly that the effect is not one of economic restriction of access (i.e. we are not viewing evidence of an enabling income effect, which would show as a positive relationship between use and income class), the income parameter is reflecting an opportunity cost effect (i.e. the professional and wage costs of being ill) or is acting as a proxy for the impact of variables that were not included in this analysis (e.g. health status, severity of disease, and sex-specific biases in frequency and duration of hospital use).

In fact, there are several important potential sources of bias in the analysis contained in this chapter, particularly in relation to the effect of family income. The first and most obvious is that, although we incor-

porated diagnostic cost differences in our benefit measures, we were unable to build in differences in severity of illness within diagnoses except insofar as length of stay is a proxy for severity. Secondly, while a number of our variables undoubtedly serve as partial proxies for health status (e.g. HAGE, REMP, and possibly INC and SCOV), it is generally recognized that the low R^2 reported by this type of micro-analysis is due primarily to the absence of sound ex ante measures of health status. To the extent that the INC variable may be correlated with health status, the parameters for this variable are likely to overstate the income-attributable differentials between income classes (if low-income families have poorer health status). Of course, that is counteracted in direction, if not magnitude, by the strong possibility that those differentials are understated as a result of exclusions in the sampling process. It was noted earlier that families with institutionalized members had a lower than average probability of ending up in our sample and that such a bias would likely be most severe in small, aged, and low-income families.

Although the model in this chapter assumes an unambiguous direction of causation, one could in fact argue convincingly that a number of purportedly independent variables are dependent variables (the case of SDOC was one example noted earlier). Finally - and again we focus on the inferred 'exogenous' role of the explanatory variables - it is not certain whether sets of those variables are without causal links. Here we have in mind the well-known ambiguity related to the possible interaction of income and education. If, for example, income is partly determined by education, then a different statistical technique should be applied. It is not clear that further statistical refinements, however, would have been justified in view of our problems of the smallness of the sample and the absence of health status variables.

6
SUMMARY AND IMPLICATIONS FOR GOVERNMENT POLICY

The principal objectives of this study have been (i) to establish and describe the patterns of hospital use and benefits by socioeconomic class in Ontario for a twelve-month period in 1974-5, (ii) to determine the relative importance of different socioeconomic characteristics in explaining the variation in hospital use and benefits across socioeconomic classes, and (iii) to introduce several different measures of hospital benefits and compare benefit incidence results using each.

In Chapter 2 we reviewed briefly the empirical studies similar in scope to (i) or (ii) above under publicly financed hospital insurance. We noted the relative dearth of such studies in the United Kingdom, the United States, and Canada. The scarcity was attributed to the difficulty in obtaining data sets that are large enough and that link hospital use with the individual's or family's socioeconomic status, health status, and so on. In addition, we identified a number of conceptual and analytical issues with which this type of study must deal. Of particular note was the crude nature of hospital benefit measures (if any) used in past studies. In fact, in view of the limitations of our sample, the method of analysis in this study seems at least as important as the results.

Chapter 3 placed considerable emphasis on the definition of the family unit used in our study, a definition unique (to our knowledge) to Ontario's health insurance plan. Among several other methodologic considerations that occupied our attention, the most notable were the construction of the data base and the conceptual and analytical issues surrounding hospital use and the measurement of benefits. We developed six different benefit measures by combining either days' stay or number of separations with various measures of per diem and per separation costs.

The empirical analysis consisted of two parts. Chapter 4 contains a conventional cross-tabulation analysis that yielded a detailed description

of patterns of hospital use in relation to three socioeconomic variables. The second part was an analysis of the importance of various largely socioeconomic family characteristics in explaining differences between families in hospital use and benefits. Part of Chapter 5 investigates empirically a number of analytic techniques that could be applied to this type of analysis. Ordinary least squares and Tobit techniques applied to our entire sample were rejected in favour of a two-stage process that used a Probit model to investigate the admissions process, and ordinary least squares regression to analyse variations in hospital use among families that received some hospital care in the study period.

The principal finding from the cross-tabulation analysis is that, although there are not significant differences between income classes in the frequency of hospitalization, the poor do make significantly greater use of, and benefit more from, insurance coverage of acute care hospitals in Ontario than other income groups, both on a per family and per capita basis. This result is primarily attributable to longer average lengths of stay, which, in turn, are a partial reflection of the high proportion of aged families in that income class. These results were maintained throughout our application of different benefit estimates, though progressive refinements to those estimates tended to reduce slightly the significance of the differences between income classes. We also found that the use and benefit differentials that do exist are for the most part a result, not of differences in diagnostic mix, but of the fact that the poor tend to be treated in relatively low-cost hospitals. The use of our various benefit measures also suggested that families with heads of 35-49 years of age were treated in relatively higher-cost hospitals, as were larger families (both of which were undoubtedly at least partly the same phenomenon). Single persons were more likely to be treated in lower-cost hospitals.

Although we found no significant difference in family frequency of hospitalization by age of OHIP family head for our sample, when attention was focused on the subsample of families with some hospital treatment, a significant positive relationship was evident between hospital use and benefits and the age of the family head. On a per family basis, larger families are, not surprisingly, more frequent users of acute-care hospitals, although the single-person family shows the greatest benefits when the denominator is the individual rather than the family.

These results generally conform to those of the studies we reviewed

in Chapter 2. In particular, our finding that there is an inverse relationship between family income and hospital use or benefits is in accordance with the findings of Statistics Canada (1977) and Boulet and Henderson (1979).

In Chapter 5, income was found not to be a significant factor underlying the probability that a family will have at least one hospitalization experience in a year. This may, however, be a result of possible exclusions from our sample of families with one or more institutionalized members at the time of sampling. Such exclusions are likely to be more prevalent among the aged and small families, which tend quite often to be one and the same.

The importance of income class shown in the cross-tabulation results re-appears when the subset of families receiving some hospital care is subjected to the ordinary least squares utilization analysis. In Figure 1 it was shown that, again, the significance of the income-class-specific benefit differentials is reduced by using hospital-specific and diagnosis-specific marginal costs per day in computing benefit estimates. The pro-poor pattern however, remains intact.

Of particular interest is the benefit measure based on MCD. The economic significance of this measure is that it does not confer or allocate the fixed cost component of hospital costs only to hospital users. These fixed costs are treated as a public good or a manifestation of 'option demand' (Weisbrod 1964) and allocated implicitly to the entire insured population rather than just to the hospitalized population, as implied to a greater or lesser degree by the other benefit measures.

Even the benefit estimates incorporating the diagnostic information are not the end of the possible methodologic refinements. It is both possible and likely that marginal costs vary not only across hospitals and diagnoses but also across age and sex categories within any hospital and diagnostic group. The method used in Barer (1981b) to generate the MCD figures for Ontario could be extended to estimate hospital-specific, diagnosis-specific, age-sex-specific marginal inpatient costs per day. These figures could be used to generate yet another set of benefit estimates, which could in turn serve as an alternative dependent variable. Perhaps a more valuable first extension, however, would be to apply the benefit measures developed so far to a richer, much larger data set such as the one used by Horne (1978) and Beck and Horne (1980). In addition, now that the

benefit incidence analyses of both the medical and hospital insurance programs have been completed, it remains to determine the tax incidence for the province by income class over the same period. Combining the three components will yield information on net fiscal incidence.

Our study has used a cross-section (static) analysis, which considered the patterns of use and benefit for the hospital insurance program in Ontario for a single year. Thus, we have not examined the changing distribution of hospital care over time. However, evidence from the Canada Sickness Survey of 1950 (i.e. for a period before the introduction of universal hospital insurance) revealed a pro-rich distribution of hospital use (Manga 1978), and evidence from the United States suggests a similar pattern. Therefore we can surmise rather tentatively that the introduction of such a scheme has increased the consumption of hospital services by the poor in relation to the non-poor, and to that extent it has been successful in fulfilling one of the principal objectives of public health insurance.

The results of our analysis suggest a number of implications for government policy. First, the earlier benefit incidence study of the medical insurance component of OHIP (Manga 1978) found that families in the second-lowest income class ($4000-7999) were the lowest users and beneficiaries of medical services. Here we have found that that income class has low use and benefits in relation to the income class below it but also in relation to the one immediately above it. Combining these results suggests that families in the $4000-7999 income class were one of the lowest, if not the lowest, income class beneficiary group in the entire Ontario Health Insurance Program. Yet a number of families in this income group are surely what one would describe as the working poor - barely above the cutoff point for exemption from OHIP premiums. That these premiums are essentially a regressive tax and are subject to a number of other criticisms has been well documented recently (Ontario Economic Council 1979, Barer, Evans, and Stoddart 1979). The regressivity of those premiums is all the more conspicuous in the light of the benefit incidence pattern of OHIP suggested by Manga's (1978) results and by the present study.

Finally our analysis also suggests that any scheme of direct charges linked to health service benefits - a variant of a tax scheme proposed by the Ontario Economic Council (1976) - will be subject to benefit measurement difficulties, apart from economic considerations that suggest that the

idea is unworthy of implementation (Barer, Evans, and Stoddart 1979). The more sophisticated hospital benefit measures are difficult to establish on a regular basis and they, too, need further refinement. Moreover, the distribution of such a tax burden would depend upon which measure was used. The results presented here suggest that the cruder the measure chosen, the more regressive (anti-poor) the burden would be, because the cruder measures have tended to overstate the hospital benefits received by the poor.

APPENDIX

ESTIMATION TECHNIQUES

Regression analysis attempts to measure the association between a dependent variable and each of several covariates. The necessary mathematical and statistical assumptions are well know, and since they all hold only rarely, vague references are often made to robustness and the central limit theorem. While some of this faith is justifiable, robustness has no clearly delineated boundaries. At some point one can no longer appeal to these 'escape' conditions. Yet that point is often vaguely defined. A full analysis would emphasize the distinction between violating the assumptions relating to parameter estimation on the one hand, and significance testing on the other. Robustness is, for example, particularly suspect when the error distribution is not two-tailed (for example, exponential distribution).

This appendix presents the various estimation options that we considered for our data set. We start with the naive application of ordinary least squares, then discuss the Tobit model and two-stage options, and finally describe the two-stage model that we adopted (Probit plus OLS).

ORDINARY LEAST SQUARES

The data set that we used, which contains a heavy concentration of zero values, violates two assumptions implicit in the application of ordinary least squares estimation - normally distributed errors and homoscedasticity. The attraction of this method, and thus the reason for the frequent appeal to robustness and the central limit theorem, is that it is straightforward to apply, easy to comprehend, and generates results that can be easily interpreted.

Results of applying this method to our full data set (n = 1290) are presented in Table A.1 for one of the four formulations of equation (3). The dependent variable is estimated hospital benefits per family, using

MCD (marginal hospital-specific and diagnosis-specific inpatient per diems) as the cost weight for computing benefits.

This analysis yielded an R^2 of 0.124 and a partial F-score on the income variables of $F_{(5,1267)} = 1.55$ ($p = 0.17$). One can have only very slight confidence in the validity of these results, however, owing to the assumptions required for the analysis.

TOBIT AND TWO-STAGE ANALYSES

The Tobit model, which was developed by Tobin (1958) to explain expenditures on luxury items, allows for a concentration of zeros and a spread of positive values. The mechanism for obtaining Tobit values assumes that observations, positive and negative, are generated from an underlying normal distribution. Negative values are then replaced by zeros (see Figure A.1). This model allows a mathematical formulation for observations that have a concentration point (in this case zero) in addition to values distributed over a continuous scale.

The Tobit model, if applied to hospital days' stay, assumes that, as μ (the mean of the underlying normal distribution) increases, there are two consequences: (1) the probability of admission will increase, and (2) the mean length of stay for those admitted will increase. This is a strong assumption. It implies that the effect of income, say, on the probability of admission will be the same as its effect on days' stay for families with admissions. As is shown below, the assumption appears untenable. A more general model would estimate separately the relationship between income and the probability of at least one admission, and then investigate the effect of income on length of stay for the subsample with admissions.

Let Z_i be a random, N (μ_i, σ^2) variable. For the Tobit model specification,

$$Y_i = 0 \text{ if } Z_i \leq 0$$
$$Y_i = Z_i \text{ if } Z_i > 0.$$

The covariates (X_i) enter the model through μ_i:

$$\mu_i = b_0 + b_1 X_1 + \ldots + b_k X_k.$$

TABLE A.1
Regression results using OLS and Tobit, with BFT_4 (see Table 18) as the dependent variable (n = 1290)

Variables	OLS coefficient	OLS T-statistic	Tobit coefficient
Constant	-187.6	-1.62	-255
INC1	54.7	0.80	79
INC2	33.7	0.55	-174
INC3	74.8	1.26	-31
INC4	4.4	0.07	-116
INC5	9.3	0.15	-214
FAM1	-59.4	-1.24	-258
FAM2	43.7	0.95	119
FAM3	122.2	2.38**	255
FAM4	-121.6	-2.54**	-644
HAGE1	262.0	5.56*	194
HAGE2	214.2	4.64*	61
HAGE3	264.6	5.78*	123
DRPOP	-9.9	-0.52	-86
BDPOP	12.9	1.29	48
HEDU1	31.6	0.94	-48
HEDU2	62.3	1.59	15
HEDU3	12.6	0.34	-119
SDOC	128.3	3.40*	132
SCOV	52.1	2.39**	129
REMP1	-106.1	-1.83	-475
REMP2	-85.4	-1.21	-394
REMP3	370.1	6.07*	217
	R^2 = 0.124		σ = 902
	For INC, $F_{(5,1267)}$ = 1.55 ($p \cong 0.17$)		For INC, $\chi^2_{(5)}$ = 6.78

* Significant at 0.01 level
** Significant at 0.05 level
Significance indications based on two-tailed tests

Let $\phi(x) = \exp(-x^2/2)/\sqrt{2\pi}$

and let $\Phi(x) = \int_{-\infty}^{x} \phi(t)\, dt$.

Then $E(Y_i) = \mu_i \Phi(\mu_i/\sigma) + \sigma\phi(\mu_i/\sigma)$,

$$\text{Prob}(Y_i > 0) = \int_{0}^{\infty} \phi((x-\mu_i)/\sigma)\, dx$$

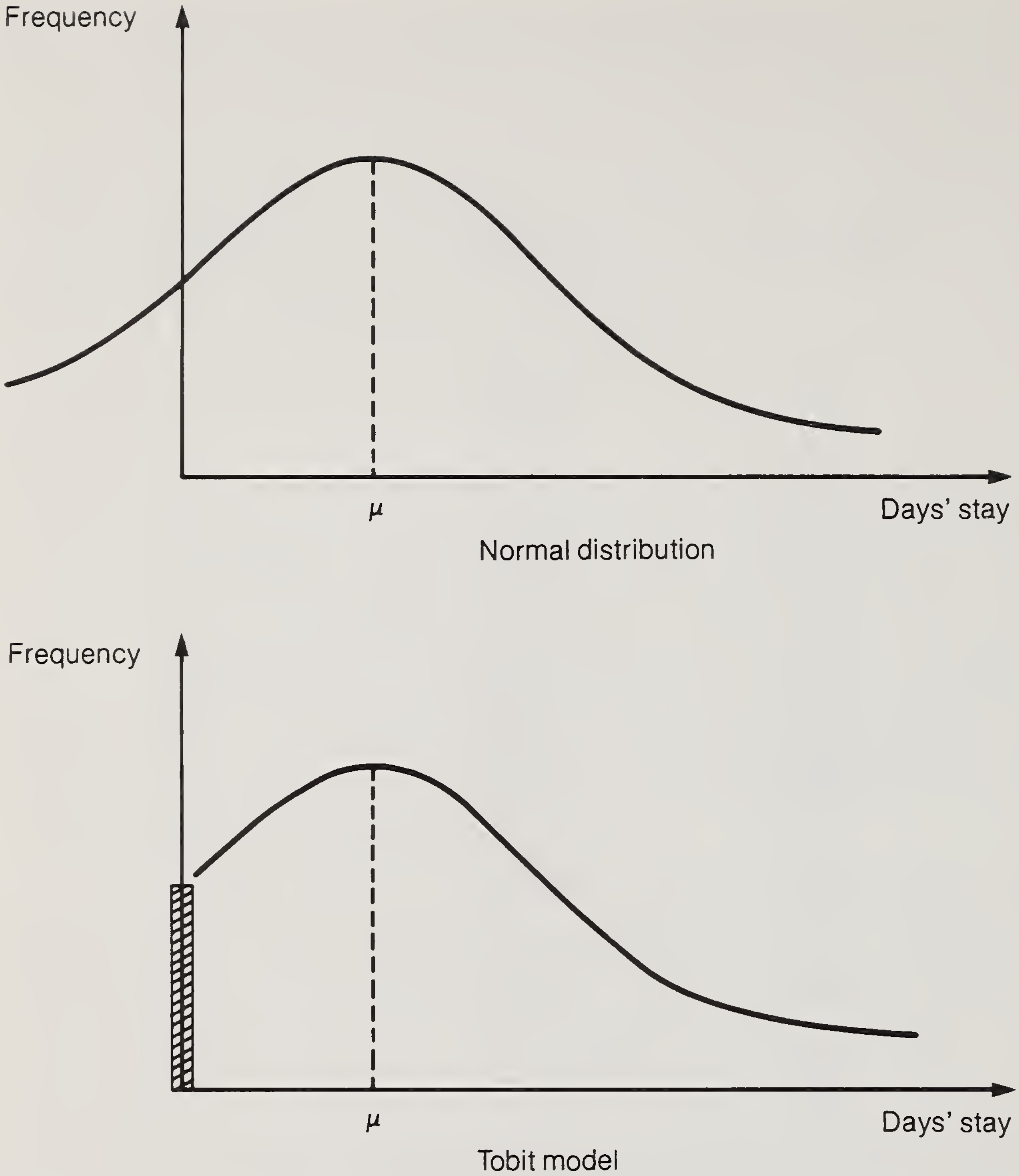

Figure A.1

$$= 1 - \Phi(-\mu_i/\sigma)$$

$$= \Phi(\mu_i/\sigma),$$

and the likelihood contribution of a Tobit observation Y_i is

$$L_i = \begin{cases} \phi((Y_i - \mu_i)/\sigma)/\sigma & \text{if } Y_i > 0 \\ \Phi(-\mu_i/\sigma) & \text{if } Y_i = 0 \end{cases} \qquad (1)$$

Thus it can be seen that as μ_i increases, the probability that $Y_i > 0$

increases. Also, the mean value of the non-zero Y_i's increases. The Tobit model assumes that the parameters of the underlying normal distribution that determine the probability of an expenditure, $\Phi(\mu_i/\sigma)$, are the same in sign and magnitude as the parameters that determine the quantity spent (for Y_i's >0).

There is no obvious reason to question this assumption when analysing expenditure on luxuries. But the situation for hospital care may be quite different. For example, younger persons may be admitted relatively frequently to hospitals as a result of accidents, but once admitted may heal relatively rapidly as a result of their youth. It is, in any event, not clear whether the Tobit model is suitable for health care, although it has been applied to it (see, for example, Davis and Reynolds 1976). The results for our data set using the Tobit estimation are reported in Table A.1. Income was not a significant variable ($\chi^2_{(5)}$ = 6.78).

The appropriateness of the Tobit model may be tested using a log-likelihood ratio test. The contribution to the likelihood from an observed Tobit Y_i was noted above in equation (1). Now, the Tobit model generalizes naturally into a two-stage model - a Probit to determine the probability of admission and a truncated normal to model the days' stay of families with at least one admission. The Probit model has likelihood

$$L_i = \begin{cases} 1 - \Phi(-\mu_i/\sigma) & \text{if } Y_i > 0 \\ \Phi(-\mu_i/\sigma) & \text{if } Y_i = 0. \end{cases} \tag{2}$$

The truncated normal is defined here to have likelihood

$$L_i = \begin{cases} \dfrac{\phi((Y_i - \mu_i)/\sigma)}{\sigma\Phi(\mu_i/\sigma)} & \text{if } Y_i > 0. \end{cases} \tag{3}$$

If the means, μ_i, in the three models are equal, then we see that the Tobit likelihood, L_i, is the product of the Probit and truncated normal likelihoods. The appropriateness of the Tobit may be tested by comparing the log-likelihoods obtained by fitting Probit and truncated normal separately.

Suppose there are s covariates, X_j, for each observation, and let MLL_T, MLL_p, and MLL_{TN} be the maximum log-likelihoods for the Tobit, Probit, and truncated normal models, respectively. Note that MLL_p has s covariates because the μ_i's cannot be estimated independently of σ. MLL_T and MLL_{TN} each have s + 1 parameters (including σ). If the Tobit model is appropriate, then the test statistic $\Delta MLL = 2(MLL_p + MLL_{TN} - MLL_T)$ is a χ^2 random variable on s degrees of freedom. This tests for the equality of the parameters in Probit and truncated normal models; if they are equal, then the Tobit is appropriate.

When the three models were fitted to the same set of data, the results were:

	Tobit	Probit	Truncated Normal
Max. LL	-3489.00	-729.60	-2675.89

The test of fit of the Tobit compares MLL = 167.02 to the percentage points of a $\chi^2_{(23)}$ (there are 23 covariates including the intercept). The difference in MLL's is an extreme value for a χ^2 distribution with twenty-three degrees of freedom. Therefore, we reject the null hypothesis, i.e. the Tobit model.

Since ΔMLL shows that the Tobit model is inappropriate, the data must be analysed in two stages. One could use separate Probit and truncated normal models, but the truncated normal is difficult to contend with from the point of view of numerical analysis. The estimation procedures frequently diverged to a point where $\Phi(x)$ could no longer be computed. Since the truncated normal only arises because of the Tobit, ordinary least squares is used in this study to analyse the non-zero values; the results based on that two-stage estimation are contained in Tables 19 and 20.

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